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Equity market linkage and multinational trade accords: The case of NAFTA

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Abstract

This paper examines the impact of multinational trade accords on the degree of stock market linkage using NAFTA as a case study. Besides liberalizing trade among the U.S., Canada and Mexico, NAFTA has also sought to strengthen linkage among stock markets of these countries. If successful, this could lessen the appeal of asset diversification across the North American region and promote a higher degree of market efficiency. We assess the possible impact of NAFTA on market linkage using cross-correlations, multivariate price cointegrating systems, speed of convergence, and generalized variance decompositions of unexpected stock returns. The evidence proves robust and consistently indicates intensified equity market linkage since the NAFTA accord. The results also suggest that interdependent goods markets in the region are a primary reason behind the stronger equity market linkage observed in the post-NAFTA period.

JEL classification: F36; F15

Keywords: NAFTA; Equity market linkage; Cointegration; Speed of convergence; Long-run price equilibrium; Generalized variance decompositions

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1. Introduction

Understanding the link between international trade and equity markets is important for pricing securities across borders, for managing multinational asset portfolios, and for global hedging strategies. This study investigates the impact of the North American Free Trade Agreement (NAFTA) on equity market linkage in the region. Ratified in January 1994, NAFTA's primary goal is to liberalize trade among the three North American countries. Evidence suggests that NAFTA has thus far succeeded in promoting trade in the region (e.g., Gould, 1996, 1998; Kouparitsas, 1997).

Although the main focus of NAFTA is on eliminating (or reducing) barriers to regional trade in goods and services, the accord has also sought to promote cross-country investment among the three economies. Article 1109, for example, calls for the free and quick transfer of all payments relating to equity transactions including dividends, interest, and capital gains. Several other provisions relax restrictions on cross-country investing and foreign stock ownership. A growing literature suggests that the evolution of equity market linkage across national borders tends to be associated with international trade and other macroeconomic factors. For example, Kouparitsas (1997) and Bracker et al. (1999) highlight the importance of interdependence among international goods markets when discussing equity market linkages. Phylaktis and Ravazzolo (2002) also report evidence suggesting that financial and economic integrations are significantly related in a group of Pacific Basin countries.

This paper examines the impact of NAFTA on the equity market linkage among the U.S., Canada and Mexico. In addition, our analysis has other important practical and theoretical implications. In particular, if stock markets in the three North American countries have become more linked in the post-NAFTA regime, then asset diversification within the region will lose much of its appeal. Moreover, a higher degree of market comovements promotes faster adjustments of equity prices to information flows from member countries, leading to more efficient markets. The degree of equity market linkage is also important for international capital budgeting. As Cooper and Kaplanis (2000) argue, optimal corporate decision rules in capital budgeting for segmented markets differ from those for fully integrated markets because similar equities provide different net returns depending on the degree of market linkages. If NAFTA has indeed strengthened equity market linkage in North America, companies will have access to capital on equal terms. This may lessen the home-country bias. Another implication pertains to possible welfare gains. Interdependent equity markets complement the goods markets as they effectively provide the price-discovery function for international trade and facilitate an optimal timing for international asset claims. Furthermore, under a stronger market linkage across borders, a multinational version of the capital asset pricing model becomes a more appropriate framework than the domestic version commonly used for individual equity markets.

Our paper contributes to the literature in several other respects. First, prior research suggests that equity market linkage has a time-varying characteristic

(Longin and Solink, 1995; Karolyi and Stulz, 1996). Bekaert and Harvey (1997) and Ng (2000) also argue that liberalized markets tend to correlate more. Yet, there is little research on the potential impact of multinational trade accords like NAFTA on equity market linkage of member countries. We attempt to fill this void in the literature.

Second, we use several econometric methods to investigate the impact of NAFTA on the degree of equity market linkage. The methods include (i) analyzing long-run equity price equilibria, (ii) examining the speed of convergence to the price equilibrium, and (iii) conducting variance decompositions on unexpected returns. These methods address different aspects of the equity market linkage. Analysis of long-run equity price equilibria and disequilibrium shocks focuses on whether national equity prices in the three North American countries move together around a common equilibrium path, and examine how quickly these markets revert to price equilibria. Having found a robust long-run price relation binding the three North American stock markets, we then use persistence profile analysis to investigate the speed of convergence to price equilibrium given a system-wide composite shock. Decomposing the variance of unexpected returns addresses the question of whether unexpected return variations in a given market can be explained by innovations (news) in other markets. Throughout all these methods, we focus on whether NAFTA has influenced market linkage in the region.

Third, we also differentiate between the confounding effects of NAFTA and the Mexican peso crisis of late 1994 on market linkage. Following Bracker et al. (1999), we model long-run linkages among equity markets as a function of several variables including those representing the goods markets disequilibrium, the NAFTA accord, and the Mexican peso crisis.

Section 2 discusses the data and their stochastic properties. Section 3 analyzes the impact of NAFTA on the long-run equity price relations. Section 4 examines the speed of the markets' adjustment to price disequilibria. Section 5 presents the empirical results for discriminating among possible factors affecting the degree of market linkage. Section 6 examines the variance decompositions of unexpected returns before and after the passage of the NAFTA accord. Section 7 concludes the paper.

2. Data and their stochastic properties

Our data on stock prices consist of the S&P 500 Composite index for the U.S., the TSE 300 Composite index for Canada, and the IPC index for Mexico. We use daily data provided by Commodity Systems Inc., spanning the period June 1, 1989 through April 10, 2002 (3040 observations). Our sample starts in June 1989 to avoid contaminating the results by the October 1987 stock market crash, or by the May 1989 market liberalization in Mexico.

Besides daily data, we also use weekly and monthly series for robustness purposes since daily data may suffer from problems associated with non-trading, non-synchronous trading, and the bid/ask spreads (Lo and MacKinlay, 1988). We convert the daily data to 639 weekly observations using Wednesday's figures.[2] The monthly data (154 observations) are end-of-month figures. Following Longin and Solnik (1995) and Bracker et al. (1999), we measure stock prices in local currencies to incorporate possible hedging activities of investors against foreign exchange risks. However, we also perform our tests using U.S. dollar denominated data to ensure the robustness of our results.

Table 1, panel A, provides summary statistics for the weekly stock returns (logarithmic first-differences of stock prices) in both local currencies and U.S. dollars for the full period, as well as for the pre- and the post-NAFTA periods. As the table indicates, the Mexican market exhibits the highest weekly average returns in local currency in all periods. However, when returns are converted to U.S. dollars, the Mexican market provides the lowest return, especially in the post-NAFTA period due, of course, to the substantial devaluation of the Mexican peso starting in December 1994. The Mexican stock market also exhibits the highest return volatility, both in local currency and in U.S. dollars, and the Canadian stock market exhibits the lowest returns and volatilities in all periods.

The cross-country return correlations, displayed in panel B of Table 1, are all statistically significant in the three periods when measured in local currencies. The magnitudes of these correlations increased in the post-NAFTA period. Furthermore, return correlations between the U.S. and Mexico, and between Canada and Mexico, are high and significant in the pre-NAFTA period when measured in local currencies. However, these correlations are not significantly different from zero when measured in U.S. dollars. Thus, return correlations among the three North American stock markets fall when returns are expressed in common currencies. This implies that the exchange-rate effect, if it exists, tends to isolate the three markets at least in the pre-NAFTA period. In the post-NAFTA period, the equity market linkage appears to have dominated the exchange-rate effect since these markets exhibit stronger and more significant correlations, even when measured in U.S. dollars. Such evidence suggests that the three North American stock markets became more linked after the passage of the NAFTA accord, while the peso crisis may have reduced the degree of this linkage. Nevertheless, we caution that contemporaneous correlations may not reflect genuine information on markets interrelationship, and that such correlations could be misleading if investors have long holding periods (Kasa, 1992; Darrat and Zhong, 2002).

We begin our analysis by subjecting the data series to a rigorous pre-testing for unit roots on the basis of several procedures (the Weighted Symmetric, the Augmented Dickey–Fuller and the Phillip–Perron tests). We search for optimal lag profiles in these tests using the Akaike Information Criterion (AIC), allowing for up to 12, 26, and 40 lags for monthly, weekly, and daily data, respectively. Since all three tests yield similar inferences, we only report the WS test results to conserve space.[3] Table 2 reports the results for monthly, weekly and daily data. These results

[2] If a Wednesday has no trading, we use the Thursday data. If that Thursday is missing for any country, we use the Tuesday's counterparts. If that Tuesday is also missing, we skip that week altogether. This approach to convert daily data to weekly data is akin to that of Lo and MacKinlay (1988).

[3] Pantula et al. (1994) argue that the WS test is more powerful than several other alternatives.

Table 1
Summary statistics of weekly stock returns

	Mean (%)						Standard deviation (%)					
	Full period		Pre-NAFTA period		Post-NAFTA period		Full period		Pre-NAFTA period		Post-NAFTA period	
	Local	US$	Local	US$	Local	US$	Local	US$	Local	US$	Local	US$
Panel A: Summary statistics												
U.S. S&P 500	0.20	0.20	0.17	0.17	0.21	0.21	2.16	2.16	1.92	1.92	2.29	2.29
Canada TSE	0.12	0.07	0.07	0.04	0.14	0.09	2.09	2.31	1.59	1.76	2.33	2.55
Mexico IPC	0.50	0.07	0.96	0.75	0.25	−0.01	4.29	5.49	3.66	3.23	4.59	5.70

Correlations between	Full period		Pre-NAFTA period		Post-NAFTA period	
	Local	US$	Local	US$	Local	US$
Panel B: Cross-correlations of stock returns						
U.S. & Canada	0.74***	0.71***	0.67***	0.62***	0.76***	0.74***
U.S. & Mexico	0.49***	0.40***	0.40***	0.05	0.52***	0.48***
Canada & Mexico	0.46***	0.42***	0.26***	0.04	0.53***	0.47***

The statistics are Wednesday-to-Wednesday continuously compounded (log) returns, both in local currencies and in U.S. dollars. The full period covers June 1, 1989 to April 10, 2002. The pre-NAFTA period covers June 1, 1989 to December 31, 1993, while the post-NAFTA period covers January 1, 1994 to April 10, 2002. The cross-correlations are contemporaneous correlations between each pair of the stock returns. *** Indicates rejection of the null hypothesis of no correlation at the 1% significance level.

indicate that all three price indexes (expressed in logs) are not stationary in levels over the full period, and also over the pre- and post-NAFTA periods. However, the null of non-stationarity is soundly rejected at the 5% level for all three price indexes in all three periods once these indexes are converted to first-differences. Thus, the three price indexes are integrated of order one.

3. Long-run equity price relations

Fig. 1 plots standardized price indexes for the three North American countries for the pre- and the post-NAFTA sub-periods.[4] Equity prices in all three countries follow an upward trend. Two issues emerge. First, equity price movements in the region have similar turning points. Does this mean that these markets share a long-run common stochastic trend, or is this simply a "spurious" association? The presence of a common stochastic trend could reduce long-term gains from portfolio diversification across countries. Hung and Cheung (1995), Darrat and Zhong (2002), and Bessler and Yang (2003) suggest that world stock markets generally exhibit common stochastic trends. One might ask: if North American stock markets share a common trend, has NAFTA introduced a regime shift in this common trend? That is, have the three North American stock markets moved more closely together as a result of NAFTA?

[4] We obtain the standardized series by subtracting the mean and then dividing by the standard deviation.

Table 2
Pantula et al.'s weighted symmetric unit root tests (p-values)

	U.S. S&P 500	Canada TSE (local currency)	Canada TSE (US dollars)	Mexico IPC (local currency)	Mexico IPC (US dollars)
Monthly data					
Levels					
Full period	0.99	0.40	0.68	0.54	0.90
Pre-NAFTA period	0.12	0.31	0.88	0.41	0.56
Post-NAFTA period	0.99	0.63	0.75	0.40	0.21
First-differences					
Full period	0.00**	0.00**	0.00**	0.00**	0.00**
Pre-NAFTA period	0.00**	0.01**	0.01**	0.01**	0.00**
Post-NAFTA period	0.00**	0.00**	0.00**	0.00**	0.00**
Weekly data					
Levels					
Full period	0.96	0.47	0.71	0.12	0.90
Pre-NAFTA period	0.10	0.84	0.98	0.99	0.66
Post-NAFTA period	0.99	0.70	0.75	0.53	0.11
First-differences					
Full period	0.00**	0.00**	0.00**	0.00**	0.00**
Pre-NAFTA period	0.00**	0.00**	0.00**	0.01**	0.00**
Post-NAFTA period	0.00**	0.00**	0.00**	0.00**	0.00**
Daily data					
Levels					
Full period	0.98	0.39	0.66	0.20	0.90
Pre-NAFTA period	0.13	0.69	0.96	0.99	0.79
Post-NAFTA period	0.99	0.63	0.54	0.54	0.09
First-differences					
Full period	0.00**	0.00**	0.00**	0.00**	0.00**
Pre-NAFTA period	0.00**	0.00**	0.00**	0.00**	0.00**
Post-NAFTA period	0.00**	0.00**	0.00**	0.00**	0.00**

This table reports the unit root test results for the log price levels and log returns of the three national stock indexes in both local currency and U.S. dollars. The monthly data are end-of-month observations. The weekly data are Wednesday-to-Wednesday observations. The full period is June 1, 1989 to April 10, 2002, the pre-NAFTA period is June 1, 1989 to December 31, 1993, and the post-NAFTA period is January 1, 1994 to April 10, 2002. The numbers reported are p-values of Pantula et al. (1994) weighted symmetric test statistics for unit roots. An intercept term and a time trend are included in the unit root tests. ** Indicates rejection of the null hypothesis of non-stationarity at the 5% significance level.

To address these important concerns, we test for the presence of long-run (cointegrating) relations among the three stock markets in the context of different regimes. These tests help assess the tendency of the markets to move together over the long term, but without ruling out short-run departures. The steady-state equilibrium relation among the three stock prices at time t in a multivariate context can be written as:

$$\sum_{i=1}^{3} w_i \left(E_{t-1} p_{i,t}\right) = e \tag{1}$$

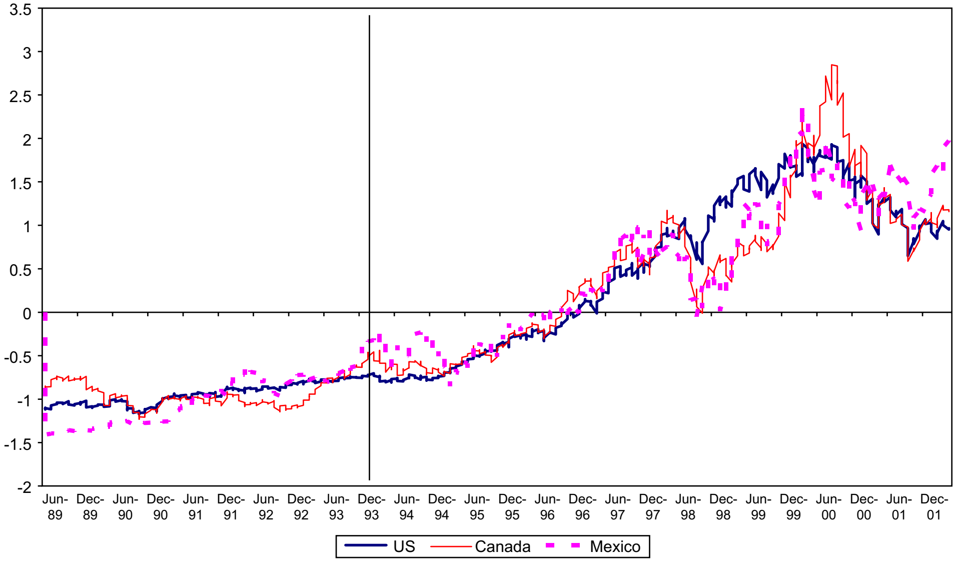

Fig. 1. Plots of standardized stock price indices of U.S., Canada, and Mexico. The vertical line in the middle of the figure corresponds to the inception of NAFTA (January 1, 1994).

where $p_{i,t}$ is the log stock price of country i in period t ($i = 1$, 2, 3, denoting U.S., Canada, and Mexico, respectively), E_{t-1} is the expectational operator, w_i is the long-run coefficient interpreted as the weight for stock market i in price equilibrium, and e represents the steady-state equilibrium across markets. Temporary deviations from equilibrium (ζ_t) are due to shocks in economic fundamentals, and are represented as:

$$\zeta_{i,t} = \sum_{i=1}^{3} w_i p_{i,t} - e. \tag{2}$$

In a matrix notation, we have:

$$\zeta_t = W' p_t, \tag{3}$$

where $W = [w_1, w_2, w_3, -e]'$, $p_t = [p_{1,t}, p_{2,t}, p_{3,t}, 1]'$, and the error term ζ_t bounces within a stationary band defined by temporary market imperfections.

The above error term can be viewed as a stationary mean-reverting component of a multinational stock price system. When a deviation from an equilibrium ensues, equity prices tend to revert back to their equilibrium values through adjustments in expected returns across countries. Such disequilibria are temporary in nature and likely the result of shocks to the fundamental values in a given market. Expected stock returns correspond to the discount rates that relate the fundamental values of stocks to their expected future dividends. Different investors' preferences regarding current versus risky future consumption, along with the stochastic evolution of firms' investment opportunities, could also cause variations in expected returns. Therefore, shocks to expected returns in one country may temporarily drive the relation of fundamental stock values across the three countries away from their long-term equilibrium. However, such differences in expected returns across countries should disappear through arbitrage, forcing equity prices to revert back to their long-term equilibrium values. If NAFTA has indeed strengthened market linkage in the region, then these markets should become more able to absorb informational shocks and revert back more rapidly to long-term equilibrium values.[5]

We first test for the presence of the long-run equilibrium price relation among the three equity markets using the Johansen and Juselius (1992) cointegration test.[6] Table 3, panels A through E, reports the cointegration test results for all three periods. To allow for the possibility of a drift in the long-run equilibrium relation, we include an intercept in the cointegrating vector. We select optimal lags in the underlying vector autoregression models (VARs) by means of Sims' Wald test, provided that the residuals are also white-noise (Cheung and Lai, 1993). The optimal lags prove to be three for monthly data, eight for weekly data, and 40 for daily data. As panel C indicates, the null hypothesis of no cointegration ($r = 0$) is strongly rejected in the post-NAFTA period using monthly, weekly and daily data. However, the hypothesis of no cointegration for the pre-NAFTA period is not rejected even at the weaker 90% level of significance across all data horizons (see panel B). Indeed, the evidence against cointegration in the pre-NAFTA period seems quite strong since it almost overturned the cointegration verdict for the full period using weekly data (see panel A). Thus, NAFTA appears to have increased the degree of market linkage in the region. Note that Mexico maintains its strong link to the region despite the government attempts to restrict capital flows in the aftermath of the peso crisis.

To further gauge the possible effect of NAFTA on the degree of cointegration among the three North American stock markets, we follow Husted (1992) and re-test for cointegration after including an intercept-dummy variable in the cointegration space. If adding the dummy variable could strengthen the cointegration finding over the full period, one may infer that the event represented by the dummy variable has contributed to a stronger market linkage.[7] The intercept-dummy variable takes the

[5] Temporary departures from market efficiency may also result in price disequilibria. Literature on behavioral finance suggests that investors may form irrational expectations about market fundamentals, which could result in short-term momentums and long-term price reversals. Odean (1999) argues that overconfident investors often place too much weight on their own private information relative to public information. Such psychological tendencies could cause stock prices to overreact to private information signals and under-react to public signals. As more public information flows into the market, prices gradually revert back to the long-term equilibrium position.

[6] We focus on the trace statistics of Johansen and Juselius (1992) in light of the Monte Carlo evidence reported in Cheung and Lai (1993). The trace statistic is computed as $-T\sum_{i=r+1}^{n}\log(1-\hat{\lambda}_i)$, where T is the number of observation, n is the number of variables in the cointegration test, r is the number of linear independent cointegrating vectors, and $\hat{\lambda}_i$ is the ith smallest squared canonical correlation in Johansen and Juselius (1992).

[7] However, results from such an approach should be interpreted with caution since the additional dummy variables are likely to be stationary. In order to maintain a conservative posture, we require the presence of more than one significant cointegrating vector to reject the null of no cointegration.

Table 3
Cointegration test results for the presence of long-run equity price equilibria among the U.S., Canada and Mexico

Null	In local currency			In U.S. dollars			Critical values	
	Monthly	Weekly	Daily	Monthly	Weekly	Daily	95%	90%
Panel A: Full period (June 1, 1989 to April 11, 2002)								
$r = 0$	32.84**	33.78*	36.06**	38.19**	30.14	38.50**	34.91	32.00
$r \leq 1$	16.07	14.37	16.51	17.34	10.01	10.49	19.96	17.85
$r \leq 2$	2.78	3.50	3.65	3.50	2.82	3.61	9.24	7.52
Panel B: Pre-NAFTA period (June 1, 1989 to December 31, 1993)								
$r = 0$	21.52	23.41	25.50	28.08	22.07	19.52	34.91	32.00
$r \leq 1$	11.99	11.35	10.57	8.33	9.51	10.59	19.96	17.85
$r \leq 2$	3.28	2.56	2.23	2.76	3.74	3.79	9.24	7.52
Panel C: Post-NAFTA period (January 1, 1994 to April 11, 2002)								
$r = 0$	35.68**	40.64**	38.18**	37.64**	32.32**	39.09**	34.91	32.00
$r \leq 1$	12.56	12.33	12.50	15.41	13.04	14.15	19.96	17.85
$r \leq 2$	3.06	4.23	3.60	5.99	4.32	5.79	9.24	7.52
Panel D: Impact of NAFTA on market linkage: the possibility of a one-time regime shift: (June 1, 1989 to April 11, 2002)								
$r = 0$	52.60*	50.96*	53.69**	53.66**	48.21	57.74**	53.12	49.65
$r \leq 1$	28.28	26.33	29.07	32.20*	26.47	29.40	34.91	32.00
$r \leq 2$	14.03	13.32	13.41	16.80	13.67	15.36	19.96	17.85
$r \leq 3$	6.42	5.54	5.73	5.62	4.96	6.18	9.24	7.52
Panel E: Impact of NAFTA on market linkage (the possibility of a continuous regime shift)								
$r = 0$	111.13**	116.27**	113.43**	100.55*	290.00**	407.87**	102.14	97.18
$r \leq 1$	75.83*	75.95*	76.24**	66.77	156.08**	161.04**	76.07	71.86
$r \leq 2$	45.35	42.97	45.19	43.03	58.90**	65.32**	53.12	49.65
$r \leq 3$	26.50	25.07	25.25	27.18	34.74*	36.66**	34.91	32.00
$r \leq 4$	11.04	11.96	11.16	15.07	13.90	15.40	19.96	17.85
$r \leq 5$	5.44	4.78	4.59	5.77	4.89	6.20	9.24	7.52

This table reports the Johansen and Juselius (1992) trace test statistics for the log prices of the three stock indexes in local currency and U.S. dollars (see footnote 6). The monthly data are month-end observations. The weekly data are Wednesday-to-Wednesday observations. The full period covers June 1, 1989 to April 10, 2002, the pre-NAFTA period covers June 1, 1989 to December 31, 1993, and the post-NAFTA period covers January 1, 1994 to April 10, 2002. We include an intercept in the cointegration vector estimation. The lag profiles in the cointegration tests are three for monthly data, eight for weekly data, and 40 for daily data. The 95% and 90% critical values are obtained from Osterwald-Lenum (1992). * Indicates rejection of the null hypothesis of no cointegration at the 10% significance level, while ** indicates rejection at the 5% significance level. Panels A, B, and C report the cointegration test results for the three stock indexes over three different sample periods. Panel D contains the cointegration test results when the systems include intercept-dummy variables that take a value of 1 for the post-NAFTA period, and 0 otherwise. Panel E contains the test results when the systems include (slope) interaction dummy variables for NAFTA.

value 1 for the post-NAFTA period, and zero otherwise. Cointegration results in panel D of Table 3 are not markedly different from those reported in panel A. To check if NAFTA has induced a continuous (rather than a single) shift, we introduce instead slope-dummy variables in the cointegration space. Results in panel E of Table 3 generally suggest the presence of two or more cointegrating vectors. Adding slope-dummy variables does strengthen the evidence of a cointegrating relation among the three markets. In particular, the test statistics for the null hypothesis of $r \leq 3$ for the weekly and daily data in U.S. dollars become significant at the higher 95% level, implying the presence of four cointegrating vectors. Thus, the NAFTA accord seems to have had a pronounced and continuous influence on market interrelationship in the region. Note also that when we include slope-dummy variables for NAFTA, there is some evidence for multiple cointegrating vectors, especially in weekly and daily data. As Dickey et al. (1991) point out, cointegrating vectors represent a set of economic restrictions imposed on the movement of the variables in the long run. Therefore, an economic system with several cointegrating vectors is stationary (stable) in more than one direction. In this sense, we may conclude that NAFTA has contributed to a more stable long-run linkage among the three North American equity markets.

Hansen and Johansen (1993) provide yet another metric to assess the stability of cointegrating vectors. They construct a likelihood ratio (LR) statistic by comparing the likelihood value obtained from each recursive sub-sample with the likelihood value computed under the restriction that the cointegrating vectors be fixed at particular constant values.[8] We plot these LR statistics in Fig. 2. Scaled by the 5% significance level, statistics that exceed the unitary band imply rejection of the null of cointegration stability. The plotted statistics show that the cointegrating parameters are sample-dependent in the pre-NAFTA period, but this apparent instability gradually disappears in the post-NAFTA period. This suggests that NAFTA has promoted a more stable linkage among the three North American markets.

4. Speed of convergence to price equilibria

Our results support the presence of a robust and stable long-run price relation binding the three North American equity markets. The evidence also shows that the

[8] The null hypothesis of sample independence of the cointegrating vector(s) is formulated as: $\beta \in \mathrm{sp}(\beta_\tau)$ $\tau = T_0,\ldots,T$, where T is the overall sample size, T_0 is the base-period sample size, β is the cointegrating vector over the whole sample, and β_τ is the cointegrating vector over sub-samples in the recursive analysis. The test statistic is:

$$\tau \sum_{h=1}^{r} \left\{ \ln\left[1 - \hat{\rho}_h(\tau)\right] - \ln\left[1 - \hat{\lambda}_h(\tau)\right] \right\} \quad \tau = T_0, \ldots, T$$

where $\hat{\rho}_h(\tau)$ and $\hat{\lambda}_h(\tau)$ are the solutions to the restricted and unrestricted eigenvalue problems. The statistic is distributed as χ^2 with two degrees of freedom.

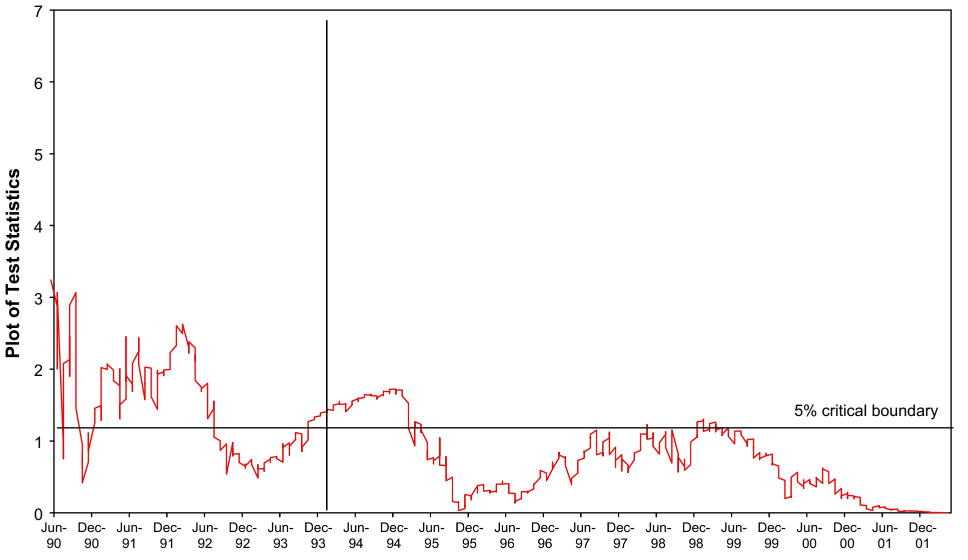

Fig. 2. Test of constancy of long-run price equilibrium. The likelihood ratio statistics assess the constancy of the long-run price equilibrium parameters using the Hansen and Johansen (1993) approach. These statistics are asymptotically χ^2-distributed and scaled by the 5% significance level, whereby values greater than 1 imply rejection of the null hypothesis of constant cointegration parameters. The vertical line in the middle of the figure corresponds to the inception of NAFTA (January 1, 1994).

cointegrating relation has gained strength in the post-NAFTA period. Therefore, North American equity markets are linked in the long run, whereby prices comove together along an equilibrium path. Should a shock to equilibrium prices occur, the markets will revert back to equilibrium. The higher the degree of market linkage, the faster the equity prices will converge. Thus, the speed of convergence can measure the degree of market linkage.

A common method to assess the speed of convergence is to compute impulse response functions (IRFs) that estimate time profiles of the effect of a shock to the cointegrated system (Lutkepohl and Reimers, 1992; Phylaktis, 1999). However, IRFs are pairwise in nature and do not allow for a system-wide response. From the market linkage perspective, the key concern is how quickly the price equilibrium relation reacts to a system-wide shock. Also, IRFs are not unique with respect to the ordering of the variables in the underlying VAR model and to the way shocks are orthogonalized.

We use persistence profile analysis to examine the time profile of the effect of a system-wide shock on the price cointegrating relation (Pesaran and Shin, 1996). The estimated profiles are unique and do not require the prior orthogonalization of shocks. This profile approaches zero as its horizon lengthens, since the effect of a shock on the cointegrating relation is transitory by nature and eventually disappears as the system reverts back to equilibrium.

We first estimate the following vector error-correction model (VECM):

$$\Delta p_t = \phi_0 + \sum_{l=1}^{\text{lag}} \phi_l \Delta p_{t-l} + NW' p_{t-1} + \varepsilon_t$$
$$\varepsilon_t \sim i.i.d.(0, \Omega), \tag{4}$$

where $p_t = [p_{1,t}, p_{2,t}, p_{3,t}]'$, ϕ_0, ϕ_l, and N are 3×3 coefficient matrices, and W is the $3 \times r$ long-run coefficient matrix (r is the number of cointegrating vectors) such that $W' p_t$ is the cointegration relation (long-run price equilibrium) as defined in Eq. (3). The resulting estimator has an asymptotic normal distribution whose covariance matrix obeys the usual formula for stationary processes. We estimate system Eq. (4) using the Johansen and Juselius (1992) method and obtain the estimates of the long-run coefficient matrix W and the covariance matrix of the residuals, Ω.

We then decompose the price vector p_t using the Beveridge and Nelson (1981) method:

$$p_t = \tau_t + \varpi_t \tag{5}$$

$$\varpi_t = \sum_{l=0}^{\infty} C_l \varepsilon_{t-l} \tag{6}$$

$$\tau_t = \mu + \tau_{t-1} + A_0 \varepsilon_t + A_1 \varepsilon_{t-1}, \tag{7}$$

$$\text{where } C_l = -\sum_{l=0}^{\infty} A_l + B_l, \text{ and } B_l = \sum_{j=0}^{l} A_j. \tag{8}$$

The price cointegrating relation can be written as:

$$\zeta_t = W' \tau_0 + (W' \mu) t + \sum_{l=0}^{\infty} W' B_l \varepsilon_{t-l}. \tag{9}$$

With one cointegrating vector, the scaled measure of the persistence profile of the cointegrating relation for the nth horizon is:

$$h_\zeta(n) = \left(W' B_n \Omega B_n' W'\right) / (W' \Omega W'), \tag{10}$$

where W and Ω are estimated by Eq. (4), and B_n is the 'cumulative effect' matrix defined in Eq. (8) and can be computed using the following recursive relation:

$$B_n = \Phi_1 B_{n-1} + \Phi_2 B_{n-2} + \cdots + \Phi_{\text{lag}} B_{n-\text{lag}}, \; n = 1, 2, \ldots, \tag{11}$$

where $B_0 = I_{3\times3}$, $B_n = 0$ for $n \le 0$, and Φ_i's come from matrices N and W in Eq. (4) using the following relation[9]:

$$NW' = I_{3\times3} - \Phi_1 - \Phi_2 - \cdots - \Phi_{\text{lag}}. \tag{12}$$

Under one cointegrating vector binding the three price series in the full as well as in the post-NAFTA periods, the persistence profiles in Eq. (10) can also be interpreted as the square of the impulse response function of the price equilibrium ζ_t to a unit composite shock u_t, where u_t is defined as[10]:

$$u_t = W'\varepsilon_t \sim iid\left(0, \sigma_u^2\right).$$

The persistence profile, following a unit composite shock to the cointegrating system, should converge to zero in a stationary system. Fig. 3 depicts the profile estimates for the price equilibrium relations in the full and post-NAFTA periods under a system-wide composite shock based on weekly data. We do not compute the persistence profile for the pre-NAFTA period since prices are not cointegrated (hence there is no convergence) over this period. We use the AIC and the restriction of white-noise residuals to select lags in the VECM (three lags prove adequate). The estimates clearly indicate that convergence to the long-run price equilibrium following a system-wide shock did gain speed in the post-NAFTA period, with about 90% of the convergence completed within 53 weeks (about 1 year) after NAFTA was enacted. In contrast, overall convergence typically exceeded 150 weeks. This implies that barriers existed before NAFTA that prevented the three equity markets from quickly converging to price equilibria. However, the NAFTA accord has apparently mitigated these barriers.

5. Linkage through interdependent goods markets or exchange-rate movements?

5.1. Cointegration analysis

The preceding cointegration analysis clearly suggests that equity market linkage has been significantly strengthened since NAFTA. We argue that this heightened

[9] The relation in (12) can be rewritten as:

$$\Phi_1 = I_{3\times3} + \phi_1$$

$$\Phi_l = \phi_l - \phi_{1-l}, \quad \text{for } l = 2, \ldots \text{lag} - 1, \text{ and}$$

$$\Phi_{\text{lag}} = -NW' - \phi_{\text{lag}-1}.$$

[10] To see this, rewrite Eq. (9) as

$$\zeta_t = W'\tau_0 + (W'\mu)t + \sum_{l=0}^{\infty} \delta_l u_{t-l}$$

with $\delta_0 = 1$, $\text{var}(u_t) = \sigma_u^2 = W'\Omega W$, and $\delta_l^2 = (W' B_l \Omega B_l' W)/(W'\Omega W) = h_\zeta(l)$.

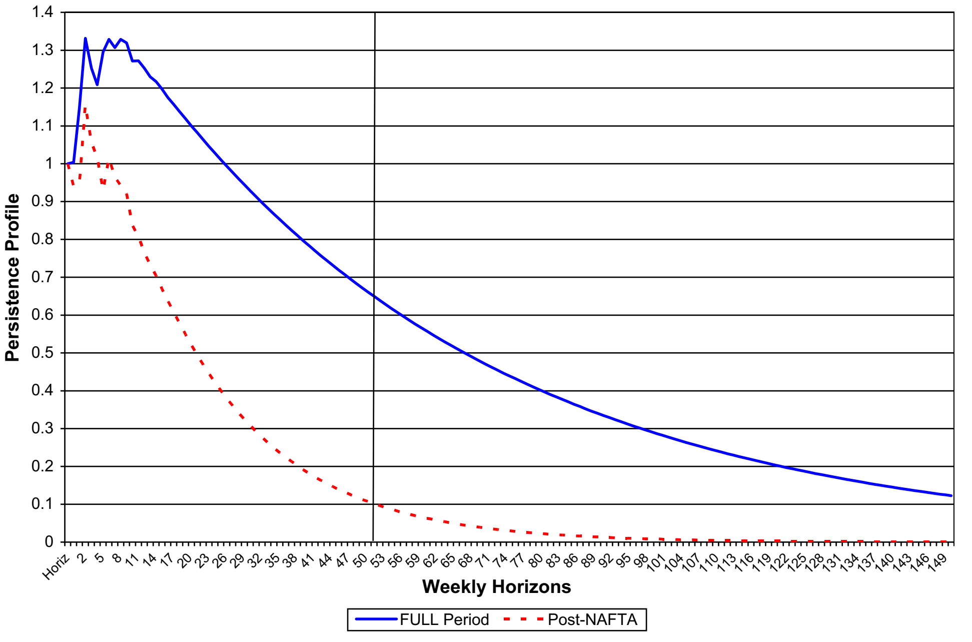

Fig. 3. Speed of convergence to price equilibrium in North American equity markets. This figure depicts the point estimates of persistence profiles based on VAR specifications for the full and post-NAFTA periods. The blue curve is the estimated persistence profile for the full period, while the red curve is the estimated persistence profile for the post-NAFTA period. The vertical line in the middle of the figure represents the 1-year (52-week) horizon.

linkage is likely the outcome of a closer goods market linkage in North America rather than due to exchange-rate movements (the Mexican peso crisis). The cointegrating relation among the industrial production indexes of the three countries, if it exists, may reflect goods market linkage. We perform the Johansen-Juselius cointegration test on the three industrial production indexes (in log).[11] Results reported in Table 4 support the presence of significant cointegration between the three indexes during the full period, as well as in the pre- and post-NAFTA periods. We also conduct cointegration tests on the two bilateral exchange rates (the Canadian dollar versus the U.S. dollar, and the Mexican peso versus the U.S. dollar). Results displayed in Table 5 indicate that the two bilateral exchange rates were cointegrated before NAFTA, but not after. Since exchange rates in the region have apparently lost their cointegration after NAFTA, it follows that the significant cointegrating relation we find among equity prices (in common US$) in the post-NAFTA period is most likely attributable to closer goods-market comovements as opposed to exchange-rate comovements.

[11] We obtain monthly data on industrial production indexes of the three countries from the *International Financial Statistics CD-ROM*. The three industrial production indexes are integrated of order one according to the unit root tests.

Table 4
Cointegration test results for the industrial production indexes of the U.S., Canada and Mexico

	Full period	Pre-NAFTA period	Post-NAFTA period	Critical values	
$r = 0$	39.54**	33.51*	38.22**	34.91	32.00
$r \leq 1$	13.30	9.91	20.76**	19.96	17.85
$r \leq 2$	3.81	1.57	4.44	9.24	7.52

This table reports the Johansen and Juselius (1992) trace test statistics for the log of industrial production indexes of U.S., Canada, and Mexico. The trace statistic is computed as $-T\sum_{i=r+1}^{n}\log(1-\hat{\lambda}_i)$, where T is the number of observation, n is the number of variables in the cointegration test, r is the number of linear independent cointegrating vectors, and $\hat{\lambda}_i$ is the ith smallest squared canonical correlation in Johansen and Juselius (1992). Monthly data are used. The full period covers June 1989 to April 2002, the pre-NAFTA period covers June 1989 to December 1993, and the post-NAFTA period covers January 1994 to April 2002. We include intercepts in the estimated cointegration vectors. The lag profiles are three in all tested cointegration vectors. The 95% and the 90% critical values are obtained from Osterwald-Lenum (1992). * Indicates rejection of the null hypothesis of no cointegration at the 10% significance level, while ** indicates rejection at the 5% significance level.

Could the stronger equity market linkage be the outcome of the Mexican peso disruption of December 1994? It is difficult to effectively discriminate between the impact of NAFTA and that of the peso crisis since the two events overlapped (the NAFTA accord was passed at the beginning of 1994, while the peso crisis erupted at the end of the same year). Nonetheless, we follow two avenues in Table 6 to shed some light on this issue. First, we perform cointegration tests on the pre- and post-peso crisis periods (see panels A and B of Table 6). We then introduce in panels C and D of Table 6 intercept and slope (interaction) dummy variables for the peso crisis to gauge its possible effects in a similar fashion to our previous analyses for the NAFTA effect (see Table 3). In contrast to our earlier evidence in support of a clear role of NAFTA in strengthening market linkage, results in Table 6 are not as supportive regarding the impact of the peso crisis. In particular, only weekly data yield significant cointegration relations at the 95% level in the post-peso crisis period (see panel B). As to monthly and daily data, the evidence for cointegration is either

Table 5
Cointegration test results for the bilateral Canadian dollar/U.S. dollar and the Mexican peso/U.S. dollar exchange rates

Null	Full period			Pre-NAFTA period			Post-NAFTA period			Critical values	
	Monthly	Weekly	Daily	Monthly	Weekly	Daily	Monthly	Weekly	Daily	95%	90%
$r = 0$	8.09	10.47	8.84	33.58**	20.66**	19.73**	17.06	12.92	15.71	19.96	17.85
$r \leq 1$	1.96	1.81	1.27	6.11	8.89	5.38	1.99	2.45	2.16	9.24	7.52

This table reports the Johansen and Juselius (1992) trace test statistics for the bilateral Canadian dollar/U.S. dollar and the Mexican peso/U.S. dollar exchange rates. The trace statistic is computed as $-T\sum_{i=r+1}^{n}\log(1-\hat{\lambda}_i)$, where T is the number of observation, n is the number of variables in the cointegration test, r is the number of linear independent cointegrating vectors, and $\hat{\lambda}_i$ is the ith smallest squared canonical correlation in Johansen and Juselius (1992). We include intercepts in the estimated cointegration vectors. The VAR lag profiles are three for monthly data, eight for weekly data, and 40 for daily data. The 95% and 90% critical values are obtained from Osterwald-Lenum (1992). ** Indicates rejection of the null hypothesis of no cointegration at the 5% significance level.

Table 6
The impact of the peso-crisis on long-run equity price equilibria among the U.S., Canada and Mexico

Null	In local currency			In U.S. dollars			Critical values	
	Monthly	Weekly	Daily	Monthly	Weekly	Daily	95%	90%
Panel A: The pre-peso crisis period (June 1, 1989 to December 31, 1994)								
$r = 0$	24.40	24.61	26.41	29.04	28.06	22.78	34.91	32.00
$r \le 1$	12.95	11.18	13.33	12.34	13.78	10.43	19.96	17.85
$r \le 2$	3.79	3.11	3.78	4.40	3.06	3.39	9.24	7.52
Panel B: The post-peso crisis period (January 1, 1995 to April 11, 2002)								
$r = 0$	32.45*	39.15**	34.03*	29.18	35.89**	32.38*	34.91	32.00
$r \le 1$	10.17	9.95	9.65	9.33	10.15	9.53	19.96	17.85
$r \le 2$	1.08	2.17	1.65	2.52	3.23	2.69	9.24	7.52
Panel C: The impact of the peso crisis on market linkage: the possibility of a one-time regime shift: (June 1, 1989 to April 11, 2002)								
$r = 0$	51.92*	51.65*	55.85**	52.16*	48.64	48.14	53.12	49.65
$r \le 1$	25.95	23.32	26.00	27.40	19.92	18.84	34.91	32.00
$r \le 2$	10.58	10.56	10.28	11.35	11.91	11.20	19.96	17.85
$r \le 3$	4.94	4.53	4.17	4.53	5.44	5.09	9.24	7.52
Panel D: The impact of the peso crisis on market linkage: the possibility of a continuous regime shift: (June 1, 1989 to April 11, 2002)								
$r = 0$	93.30	102.23**	104.69**	97.33*	121.43**	113.10**	102.14	97.18
$r \le 1$	53.25	55.21	60.00	60.44	71.70	77.15*	76.07	71.86
$r \le 2$	33.24	33.69	34.45	37.45	40.80	47.56	53.12	49.65
$r \le 3$	17.74	18.90	18.78	20.30	21.45	24.15	34.91	32.00
$r \le 4$	6.82	9.60	7.37	8.11	12.36	12.60	19.96	17.85
$r \le 5$	1.76	3.70	2.43	3.11	5.30	3.80	9.24	7.52

This table reports the Johansen and Juselius (1992) trace test statistics for the log prices of the three stock indexes. The trace statistic is computed as $-T\sum_{i=r+1}^{n}\log\left(1-\hat{\lambda}_i\right)$, where T is the number of observation, n is the number of variables in the cointegration test, r is the number of linear independent cointegrating vectors, and $\hat{\lambda}_i$ is the ith smallest squared canonical correlation in Johansen and Juselius (1992). The monthly data are month-end observations. The weekly data are Wednesday-to-Wednesday observations. The pre-peso crisis regime covers the period before January 1, 1995, while the post-peso crisis regime covers the period thereafter. We include intercepts in the estimated cointegration vectors. The VAR lag profiles are three for monthly data, eight for weekly data, and 40 for daily data. The 95% and 90% critical values are obtained from Osterwald-Lenum (1992). * Indicates rejection of the null hypothesis of no cointegration at the 10% significance level, while ** indicates rejection at the 5% significance level. Panels A, B, and C report the cointegration test results for the three stock indexes over three different sample periods. Panel D contains the cointegration test results when the systems include intercept-dummy variables that take a value of 1 for the post-peso crisis regime, and 0 otherwise. Panel E contains the test results when the systems include (slope) interaction dummy variables for the peso crisis.

dilute or non-existent. In addition, unlike the effect of NAFTA, panel D of Table 6 indicates that incorporating slope-dummy variables for a possible continuous shift resulting from the peso crisis only mildly improves the significance levels of the tests. Also unlike the case of NAFTA, adding slope-dummy variables does not lead to the emergence of several significant cointegrating vectors. Taken together, the results in Table 6 suggest that the peso crisis is not likely the main propagator for the increased equity market linkage in the region.

5.2. Regression analysis

In this section, we follow the procedure outlined by Bracker et al. (1999) to separate the effects of goods markets interdependence from the effects of the peso crisis. Specifically, we first construct dynamic measures of the equity market linkage as well as the goods market linkage. We use the resultant residuals (error-correction terms) from the significant cointegrating relation among equity prices and industrial production indexes to represent, respectively, deviations from their equilibrium paths. That is, we estimate the following model:

$$\begin{aligned}\mathrm{EC}_t = {} & b_0 + b_1\mathrm{EC}_{t-1} + b_2\mathrm{ECIP}_t + b_3\mathrm{DXCA}_t + b_4\mathrm{DXMX}_t + b_5\mathrm{NAFTA}_t \\ & + b_6\mathrm{CRISIS}_t + b_7\mathrm{NAFTA}*\mathrm{ECIP}_t + b_8\mathrm{NAFTA}*\mathrm{DXCA}_t \\ & + b_9\mathrm{NAFTA}*\mathrm{DXMX}_t, \end{aligned} \tag{13}$$

where EC_t is the error-correction term derived from the cointegration relation among the three stock market indexes (either in local currency or in US$) for the full period; ECIP_t is the error-correction term from the cointegration relation among the three industrial production indexes estimated from the full period; DXCA_t is the percentage change in the Canadian$/US$ bilateral exchange rate; DXMX_t is the percentage change in the peso/US$ bilateral exchange rate; NAFTA_t is a dummy variable taking 1 after January 1, 1994, and 0 otherwise; CRISIS_t is a dummy variable taking the value 1 after January 1, 1995, and 0 otherwise; $\mathrm{NAFTA}*\mathrm{ECIP}_t$, $\mathrm{NAFTA}*\mathrm{DXCA}_t$, and $\mathrm{NAFTA}*\mathrm{DXMX}_t$ are interaction terms with the NAFTA dummy variable.

All variables in the above model prove stationary according to several unit root tests. Table 7 reports the estimates from alternative versions of the above model which primarily assess the impact of the goods market linkage and that of exchange-rate movements. Model (1) incorporates the effect of the goods market linkage (ECIP_t) only, while Model (2) considers the effect of exchange-rate movements only. Model (3) allows for both effects, and Model (4) adds the interaction terms of these variables with the NAFTA dummy. Whether in local currency or in U.S. dollars, deviations from the long-run equilibrium path of equity prices are significantly determined by deviations from the long-run equilibrium path of industrial production indexes (ECIP_t), as well as from the NAFTA dummy variable. The positive sign on the ECIP_t variable indicates that deviations from goods markets comovement tend to be associated with deviations from the equity markets comovement. Thus, as the goods markets in the region become closely interdependent, equity market linkage will also strengthen. On the other hand, the negative sign of the NAFTA dummy suggests that the three equity markets have experienced less deviation from the long-run equilibrium path after NAFTA. The interaction term between ECIP_t and the NAFTA dummy variables is also highly significant, supporting the important dual effect of NAFTA and the goods market linkage on promoting a more potent and faster convergence to a long-run equilibrium path in the region.

Table 7
Regressions results on the equity price disequilibrium (Eq. 13)

	In local currency				In U.S. dollars			
	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4
b_0	0.018 (1.53)	0.012 (0.98)	0.013 (0.98)	0.002 (0.14)	0.012 (0.71)	0.005 (0.30)	0.004 (0.25)	−0.038** (−2.22)
b_1	0.914** (31.20)	0.945** (31.11)	0.931** (26.29)	0.938** (26.79)	0.886** (42.17)	0.933** (48.12)	0.918** (38.18)	0.923** (36.57)
b_2	0.270** (2.82)		0.184** (2.02)	0.770** (2.99)	0.688** (2.75)		0.349 (1.49)	2.466** (5.59)
b_3		0.096 (0.58)	0.094 (0.63)	−0.386 (−1.58)		0.357 (0.68)	0.314 (0.62)	0.051 (0.08)
b_4		0.069** (2.75)	0.065** (2.60)	0.021 (0.23)		0.212** (3.20)	0.205** (3.04)	0.418 (1.45)
b_5	−0.021** (−2.21)	−0.019** (−2.16)	−0.020** (−1.97)	−0.010 (−0.97)	−0.032** (−2.39)	−0.029* (−1.90)	−0.032** (−2.00)	0.012 (0.69)
b_6	0.005 (0.73)	0.007 (1.08)	0.006 (0.71)	0.007 (0.81)	0.030* (1.80)	0.025 (1.35)	0.026 (1.34)	0.027 (1.51)
b_7				−0.654** (−2.29)				−2.348** (−4.33)
b_8				0.632** (2.35)				0.386 (0.43)
b_9				0.044 (0.43)				−0.211 (−0.65)
$\overline{R}^2$	0.92	0.92	0.92	0.92	0.85	0.88	0.88	0.88

All estimates are serial correlation-consistent. The numbers in parentheses beneath the estimated coefficients are the corresponding t-values. $\overline{R}^2$ is adjusted R-squared of the regression. * Indicates significance at the 10% significance level, while an ** indicates significance at the 5% significance level.

Observe that the CRISIS dummy is not highly significant in any regression. On the other hand, changes in the Mexican peso/U.S. dollar exchange-rate exhibit statistical significance, though they appear with a positive sign when the NAFTA slope-dummy variables are excluded from the regressions. This suggests that when the Mexican peso began to fall relative to the U.S. dollar, equity markets in the region began to experience further deviations from equilibrium. Therefore, consistent with Bracker et al. (1999), exchange-rate movements exert a negative effect on equity market linkage. However, when the NAFTA slope dummy is included in the regressions, the Mexican peso loses much of its statistical weight. This suggests that the positive role of NAFTA in strengthening stock market linkage in the region dominates the adverse impact of the peso crisis on market linkage.[12]

[12] We conduct other tests using alternative dummy variables that represent the Asian currency crisis and the Russian debt default. None of these variables displays statistical significance on the degree of linkage among the North American equity markets. These additional results are available upon request.

To summarize, then, the regression results suggest that equity market linkage have intensified after NAFTA. The driving force behind this outcome appears to be the closer linkage among the goods markets in the region rather than movements in the exchange rates. This finding concurs with Phylaktis and Ravazzolo (2002) who conclude that economic integration is associated with financial integration. Similar to Forbes and Rigobon (2002), we find no evidence to support the notion that the Mexican peso crisis has significantly influenced the degree of equity market linkage in the region.

6. Variance decompositions of unexpected equity returns

Although significant cointegration among equity prices in the post-NAFTA period implies a higher degree of market linkage, it remains useful to further analyze expected returns in these markets. Since expected returns are not observable, we forecast them from a VAR process of observed (actual) returns. The resultant forecast errors represent unexpected returns (shocks). Under market segmentation, stock returns in a given market do not respond to innovations (news) from other markets. However, as market linkage strengthens, return variations in one market that are explained by their own domestic innovations should diminish, while innovations from other markets assume more explanatory power.

The unexpected returns in the pre- and post-NAFTA periods are computed as the forecast errors from a vector autoregressive (VAR) model of returns[13]:

$$r_{t+k}=\kappa+\sum_{l=1}^{k}\Theta_l r_{t+l-1}+\varepsilon_t \tag{14}$$

where r_t is the vector of continuously compounded stock returns of the three equity markets in period t [i.e., $r_t=(r_{1t},r_{2t},r_{3t})'$, where 1 = U.S., 2 = Canada, 3 = Mexico]; κ is the intercept vector; Θ_l's are autoregressive parameter matrices; and ε_t is a white-noise residual vector. The n-period forecast error of returns can be written as:

$$r_{t+n}-E_t r_{t+n}=\sum_{l=0}^{n-1}D_l\varepsilon_{t+n-l}. \tag{15}$$

The Choleski decomposition is typically used to obtain the proportions of the n-step ahead forecast error variance of variable i which is accounted for by innovations in variable j. However, this method is sensitive to the ordering of variables in the VAR. To overcome this problem, Pesaran and Shin (1998) propose a generalized

[13] Based on our earlier cointegration findings, we use a standard VAR for the pre-NAFTA period, but use a vector error-correction model (VECM) for the post-NAFTA period to capture the disequilibrium correction process (see Granger, 1986). The VECM representation incorporates the lagged error-correction terms to the VAR in Eq. (14) as an exogenous variable.

forecast error variance decomposition (GVDs) method. Specifically, the GVDs of the n-step forecast error of return i explained by return innovation j take the form:

$$\mathrm{GVD}_{ij}(n)=\frac{\sum_{l=0}^{n}\left(e_i' D_l \Sigma e_j\right)^2}{\sigma_{ii}\left[\sum_{l=0}^{n}\left(e_i' D_l \Sigma D_l' e_i\right)\right]}, \tag{16}$$

where e_i is an 3×1 selection vector with unity as its ith element and zeros elsewhere, Σ is the sum of squared of residuals, $E(\nu_t \nu_t')$, and σ_{ii} is the residual variance in the ith equation in the VAR.

Tables 8 and 9 report the GVDs point estimates of the unexpected returns (with weekly data in both local currencies and U.S. dollars, respectively) over horizons 1–5, 10, and 20 weeks, for the pre- and post-NAFTA periods. Results from monthly and daily data are qualitatively similar and are available upon request. We also compute the associated t-statistics based on Monte Carlo simulations with 500 random draws to test the null hypothesis that there is no difference in the estimated GVDs between the two sub-periods.[14,15] We can easily reject the above null hypothesis for all forecast horizons at the 5% level of significance. Specifically, whether in local currency (Table 8) or in U.S. dollars (Table 9), the percentages of the GVDs in any market that are explained by their own return innovations have significantly fallen in the post-NAFTA period (see the corresponding columns of negative differences in both tables for the U.S. in panel A, for Canada in panel B, and for Mexico in panel C). A similar message emerges from cross-variation accountings, suggesting that NAFTA has strengthened market linkage in the region. For example, the contributions of the Canadian and Mexican innovations in explaining stock returns in the U.S. market have significantly increased after NAFTA (see the corresponding columns of positive differences for Canada and Mexico in panel A of Table 8). Similarly, for the Mexican market, innovations from the U.S. and Canadian markets have claimed more explanatory power after the passage of NAFTA (see the corresponding columns of positive differences for the U.S. and Canada in panel C of Table 8). Taken together, these results further imply that the three equity markets have become more closely linked following NAFTA. As Koutmos (1997) points out, greater stock markets interdependence could be the outcome of larger capital flows in the region following the passage of NAFTA.[16]

[14] We perform the simulations first by obtaining the residual matrix from the VAR (14), and then by looping over 500 random draws to compute the corresponding standard errors of the GVDs.

[15] We test the null using the following two-tailed t-test statistic:

$$t=\frac{\mathrm{GVD}_{\text{pre-NAFTA}}-\mathrm{GVD}_{\text{post-NAFTA}}}{\sqrt{\frac{1}{\text{draws}}\left[\mathrm{var}\left(\mathrm{GVD}_{\text{pre-NAFTA}}\right)+\mathrm{var}\left(\mathrm{GVD}_{\text{post-NAFTA}}\right)+2\,\mathrm{cov}\left(\mathrm{GVD}_{\text{pre-NAFTA}},\mathrm{GVD}_{\text{post-NAFTA}}\right)\right]}}$$

where variances and covariances are computed from Monte Carlo simulations with 500 random draws.

[16] Besides first moments of returns, we have also examined market linkage through return second moments (conditional volatilities). The results, available upon request, corroborate our evidence showing increased market linkage and indicating that volatility spillovers among the three markets have intensified after NAFTA.

Table 8
Generalized variance decompositions (GVDs) of unexpected stock returns (weekly data in local currency)

Horizons	U.S.			Canada			Mexico		
	Pre-NAFTA	Post-NAFTA	Difference	Pre-NAFTA	Post-NAFTA	Difference	Pre-NAFTA	Post-NAFTA	Difference
A. Unexpected return variance of the U.S. equity returns explained by innovations in the market of									
1	59.78	53.42	−6.37**	29.74	31.70	1.96**	10.48	14.88	4.41**
2	59.44	53.54	−5.90**	30.26	31.60	1.34**	10.30	14.85	4.55**
3	58.85	53.35	−5.50**	30.13	31.84	1.71**	11.01	14.81	3.80**
4	58.85	53.33	−5.52**	30.13	31.83	1.70**	11.02	14.84	3.83**
5	58.08	53.11	−4.96**	30.14	32.10	1.96**	11.78	14.78	3.00**
10	57.79	52.99	−4.80**	30.09	32.39	2.30**	12.12	14.63	2.50**
20	57.64	52.95	−4.69**	30.21	32.38	2.17**	12.15	14.67	2.52**
B. Unexpected return variance of the Canadian equity returns explained by innovations in the market of									
1	31.77	31.73	−0.03	63.85	53.47	−10.38**	4.38	14.80	10.42**
2	32.26	31.71	−0.56	62.70	53.50	−9.20**	5.04	14.80	9.76**
3	32.05	31.71	−0.34	62.40	53.45	−8.95**	5.55	14.84	9.29**
4	32.02	31.76	−0.26	62.42	53.42	−9.00**	5.56	14.81	9.26**
5	32.01	31.67	−0.34	62.32	53.56	−8.75**	5.67	14.76	9.09**
10	31.84	31.93	0.09	61.58	53.19	−8.39**	6.58	14.88	8.30**
20	31.81	31.93	0.11	61.60	53.18	−8.42**	6.59	14.89	8.30**
C. Unexpected return variance of the Mexican equity returns explained by innovations in the market of									
1	14.09	17.91	3.83**	5.52	17.80	12.28**	80.39	64.29	−16.10**
2	14.53	18.02	3.49**	5.38	17.77	12.39**	80.09	64.21	−15.88**
3	14.45	18.16	3.70**	5.93	17.75	11.82**	79.62	64.09	−15.53**
4	14.99	18.12	3.13**	6.85	17.81	10.96**	78.16	64.07	−14.09**
5	15.46	18.21	2.75**	6.85	18.33	11.48**	77.69	63.46	−14.22**
10	16.46	18.48	2.03**	8.83	18.51	9.69**	74.72	63.01	−11.71**
20	16.45	18.50	2.05**	9.06	18.53	9.46**	74.49	62.98	−11.51**

Standard VARs are used to produce the variances of unexpected returns for the pre-NAFTA period, and, under co-integration, VECMs are used to produce the variances of unexpected returns for the post-NAFTA period. All point estimates of the GVDs are significantly different from zero since their standard errors are less than half of the point estimates. The null hypothesis that there is no difference between the GVD pointestimates before and after the NAFTA is tested using a two-tailed t-test statistic. ** indicates rejection at the 5% significance level.

Table 9
Generalized variance decompositions (GVDs) of unexpected stock returns (weekly data in U.S. dollars)

Horizons	U.S.			Canada			Mexico		
	Pre-NAFTA	Post-NAFTA	Difference	Pre-NAFTA	Post-NAFTA	Difference	Pre-NAFTA	Post-NAFTA	Difference
A. Unexpected return variance of the U.S. equity returns explained by innovations in the market of									
1	83.28	55.95	−27.32**	0.49	30.92	30.43**	16.23	13.13	−3.11**
2	87.11	56.07	−31.05**	0.49	30.83	30.34**	12.39	13.10	0.71
3	87.05	55.99	−31.06**	0.87	30.94	30.07**	12.08	13.07	0.99**
4	86.67	55.88	−30.78**	1.23	30.88	29.65**	12.10	13.23	1.13**
5	82.03	55.68	−26.35**	5.67	31.13	25.46**	12.30	13.19	0.89*
10	75.90	55.44	−20.46**	9.73	31.45	21.72**	14.37	13.11	−1.26**
20	72.01	55.42	−16.59**	12.78	31.45	18.67**	15.21	13.13	−2.08**
B. Unexpected return variance of the Canadian equity returns explained by innovations in the market of									
1	0.57	31.24	30.67**	97.18	56.52	−40.66**	2.25	12.24	9.99**
2	17.32	31.20	13.89**	80.74	56.54	−24.20**	1.94	12.26	10.31**
3	16.75	31.15	14.40**	79.52	56.45	−23.07**	3.74	12.40	8.67**
4	16.58	31.21	14.63**	79.72	56.41	−23.31**	3.71	12.38	8.67**
5	17.62	31.09	13.46**	77.76	56.57	−21.19**	4.62	12.34	7.73**
10	18.97	31.26	12.29**	73.08	56.32	−16.76**	7.94	12.42	4.47**
20	18.82	31.26	12.43**	72.74	56.31	−16.44**	8.43	12.44	4.00**
C. Unexpected return variance of the Mexican equity returns explained by innovations in the market of									
1	16.01	16.17	0.16	1.90	14.92	13.03**	82.10	68.91	−13.19**
2	15.15	16.45	1.30**	2.03	14.86	12.83**	82.81	68.69	−14.13**
3	14.24	16.44	2.21**	1.90	14.87	12.97**	83.86	68.69	−15.17**
4	13.85	16.43	2.58**	3.77	14.89	11.12**	82.38	68.68	−13.71**
5	13.97	16.51	2.54**	4.50	15.47	10.97**	81.53	68.02	−13.51**
10	17.80	16.82	−0.98**	4.89	15.46	10.57**	77.31	67.72	−9.59**

Standard VARs are used to produce the variances of unexpected returns for the pre-NAFTA period, and, under co-integration, VECMs are used to produce the variances of unexpected returns for the post-NAFTA period. All point estimates of the GVDs are significantly different from zero since their standard errors are less than half of the point estimates. The null hypothesis that there is no difference between the GVD pointestimates before and after the NAFTA is tested using a two-tailed t-test statistic. * Indicates rejection of the null hypothesis of unit root at the 10% significance level, and ** indicates rejection at the 5% significance level.

7. Conclusion

The main purpose of this paper is to examine the impact of NAFTA on the equity market linkage among the three North American countries. Since its inception in January 1994, NAFTA has sought to strengthen market linkage in the region which, if successful, could lessen the appeal of asset diversification and promote market efficiency in the North American region. We use several empirical procedures to measure the degree of market linkage, including analysis of long-run equity price equilibrium, examination of the speed of convergence to price equilibrium, and the computation of generalized variance decompositions of unexpected returns. We also perform regression analyses to examine how market linkage has evolved in the three North American countries after NAFTA.

The empirical evidence unambiguously suggests that NAFTA has indeed intensified market linkage in the region. This inference proves quite robust and persists over several tests of market linkage. The results indicate that NAFTA has promoted a significant and a more stable long-run linkage among the three North American stock markets. The presence of a stronger market linkage in the post-NAFTA period casts some doubt on the extent of benefits from international diversifications. Moreover, regression results derived from long-run equity price disequilibria suggest that one driving force behind stronger equity market linkage is likely the increased interdependence among the goods markets in the region rather than the Mexican peso crisis. These findings corroborate results reported by Bracker et al. (1999) and Phylaktis and Ravazzolo (2002) that show that linkage among equity markets may depend on the degree of economic integration. Finally, our results support recent trends in the literature to internationalize the traditional capital asset pricing model and confirm the fruitfulness of such efforts (e.g., Bekaert and Harvey, 1995; De Santis and Imrohoroglu, 1997). One possibly interesting extension of these important studies is to incorporate the underlying long-run price equilibrium when modeling international asset prices.

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