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# COINTEGRATION OF EQUITY MARKETS IN THREE COUNTRY GROUPS OF OECD COUNTRIES

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**Abstract**

This paper investigates the stock market integration in North America, the Eurozone, and other OECD countries. The main questions addressed include whether the stock markets in each of the three country groups are cointegrated and whether the relationships are stable over time. The linkages are modeled using different statistical tools, including correlation, cointegration, vector-error correction, and Granger causality tests. The results show that the correlation coefficients are noticeably higher for the two country groups that are members of a trade or currency zone (NAFTA and the Eurozone) compared to the other OECD countries. Considering full samples, the results from the cointegration tests point to no cointegration for NAFTA markets, while they are mixed in the case of the other two country groups. After controlling for structural breaks by estimating the models for different sub-periods, the verdicts from the two cointegration tests were unanimous as the results unveil evidence of cointegration in each sub-periods for each of the three country groups. The results support the hypothesis that financial markets in the OECD countries are cointegrated. They also confirm the existence of structural breaks, the dynamic nature of the long-run integration of equity markets, and the argument that markets exhibit time-varying long-run interdependence.

**Keywords**: Cointegration, Stock Market Integration, NAFTA, Eurozone, OECD, Granger Causality

**JEL Classifications:** F36, F37, G15

## 1. Introduction

Integration of world financial markets has increased, providing opportunities for efficient allocation of capital resources, better corporate governance, higher investment and saving, and galvanized economic growth (Gupta, and Guidi, 2012; Levine, 1997; Kose *et al.* 2006; Obstfeld, 1994; Van Wincoop, 1999; Eichengreen and Mussa,1998). This global market integration, however, embodies the potential for adverse effects as higher degree of financial linkages can generate a severe cross-border financial contagion amongst the integrated economies, including trade diversion, loss of control or ineffectiveness of fiscal and monetary policy by countries, and the inability for investors to decrease the overall risk of their international portfolios as they diversify internationally for better returns. This inability to reap the full benefits of global diversification derives from the higher stock return correlations brought about by the increased world financial market integration.

Convergence and co-movements of regional and world national stock markets are the result of different factors, including globalization, trade and financial liberalization, regional trade agreements, formation of currency areas, and advancements in information technology. For Chambet and Gibson (2008) and Pretorious (2002), more liberal international trade policy generates strong interdependence among stock markets and speeds up the integration and co-movement process of those markets. In the same vein, according to Ciner (2006), it can be argued that stronger economic linkages are likely to lead to increased co-movements among the national markets. Chen and Zhang (1997) and Forbes and Chinn (2004) posit that financial markets in countries with trade flows and strong economic ties are bound to move together, therefore supporting the evidence provided in the literature on the impact of international trade arrangements in reducing national equity market segmentations.

Integration and co-movements of financial markets in regional trading blocks and monetary unions continue to ignite serious debates due to the importance of their impacts on those markets and regions.[1] These and similar regional economic blocks are a clear indication of their proliferation and the emergent trend toward integrated financial markets with important implications for effective global portfolio diversification strategies. More synchronized or integrated stock markets will have serious implications on global portfolio risks, as economic theory and empirical evidence suggest that benefits from international diversification may not materialize if national equity markets are co-integrated or move together in the long run.

Although much research (Lingaraja *et al.* 2015; Bekaert and Harvey, 1995; Bertero and Mayer, 1990; Khan, 2011; Valadkhani and Chancharat, 2008) has been devoted to co-movements of the world's national stock markets, this research largely concentrates on large and non-regional markets. In turn, the impacts of regional trade and currency agreements on equity market integration have not received sufficient attention. In studies that did investigate cointegration of equity markets in regional trade blocks (Aggarwal and Kyaw, 2005; Kim *et al.* 2005; Ratanapakorn and Sharma, 2002; Maghyereh, 2006; Berben and Jansen, 2009; Lopez-Herrera and Ortz, 2010), the findings are largely inconclusive, with this issue being generally attributed to differences in methodology, model, data, and the sample and time period of study.

The current study provides additional evidence on the integration of financial markets by performing in-depth analyses of equity markets. Our examination of cointegration and interdependence of stock markets considers different issues. First, using two cointegration techniques, we examine the long-run relationship amongst markets in three different country groups, North American Free Trade Agreement (NAFTA), the Eurozone (both members of a trade or currency union), and a selected Organization for Economic Cooperation and Development (OECD) country group (not members of a trade or currency union). Second, considering the likelihood of the existence of structural breaks in the data (an issue generally ignored in previous studies) and to avoid erroneous conclusions based on tests using the full data sample, the long-run relationships are re-estimated for several sub-periods using the same two cointegration techniques. Next, we evaluate the short-run dynamic relationships amongst the NAFTA markets using the vector-error correction model (VECM) and the Granger causality test.

The paper is structured as follows. A review of the literature on international equity markets integration is presented in section 2. Section 3 explains the methodology and model of the investigation. In section 4, the empirical results are disclosed and analyzed, and section 5 summarizes the findings.

[1] Examples of regional trading blocks and monetary unions are the Association of Southeast Asian Nations (ASEAN), the Economic Community of West African States (ECOWAS), the European Monetary Union (EMU), the European Union (EU), the North American Free Trade Agreement (NAFTA), and the Southern Common Market (MERCOSUR).

## 2. Literature review

Investing in diversified portfolios of securities to increase returns and, more importantly, to reduce exposure to risk is the golden rule in modern portfolio theory. Fund managers, therefore, tend to favor international portfolios over their domestic counterparts. The global trend towards liberalization of capital flows and the growing interdependence of international equity markets, however, dampen the risk-reduction benefits of global portfolios. Given the importance and implications of equity markets interdependence for investors as well as policymakers, much attention has been devoted to financial market integration to establish its trend, degree, and dynamics (Kasa, 1992; Rangvid, 2001, Aggarwal and Kyaw, 2005; Dorodnykh, 2014, Medewitz *et al.* 1991, Menezes, 2013; Berben and Jansen, 2005). Past research on equity markets integration yielded different results, leaving our understanding of the problem still blurred. While some studies provided evidence of strong integration and co-movements of the world equity markets, others found that those markets were still segmented, at least during certain periods and in some regions of the world.

While some studies investigated the correlation coefficients between the returns of assets (Meric and Meric, 1989; Ammer and Mei, 1996; Chiang *et al.* 2007), others resorted to the existence of common stochastic trends to examine long-run cointegration among the equity markets (Aggarwal and Kyaw, 2005; Fraser and Oyefeso, 2005; Kasa, 1992; Menezes, 2013; and Click and Plummer, 2005). Cointegration among equity markets gives the evidence that markets follow the same long-run stochastic path and, as a consequence, gains from diversification across those markets may not materialize (Manning, 2002; Kanas, 1998; Ciner, 2006; Phengpis and Swanson, 2006). Aggarwal and Kyaw (2005), Chukwuogor and Kasibhatla (2007), and Benli *et al.* (2012) provided evidence of a common trend, namely the presence of long-run equilibrium relationships in groups of countries. Kasa (1992) applied the Johansen's multivariate cointegration method to investigate the permanent and transitory components of stock price series and examined the existence of a single common stochastic trend in the equity markets of the USA, Japan, England, Germany, and Canada. The study provided evidence of the presence of a single common trend driving those countries' stock markets. Chen *et al.* (2002) investigated the dynamic interdependence of major stock markets in Latin America based on cointegration analysis and concluded that there is one cointegrating vector that appears to explain the dependencies in prices. Khan (2011) examined the long-run convergence of the United States and 22 other developed and developing country markets using daily data and the Johansen (1988) and Gregory and Hansen (1996) tests. The research showed that stock markets of most countries have become cointegrated by 2010. In the same vein, Yunus (2013) investigated the dynamic interdependence among ten major equity markets throughout North America, Europe, Latin America, and Asia. Using a recursive cointegration technique, a method able to pinpoint and capture the approximate timing of a major global crisis, the study concluded that the international equity markets are integrated and that the degree of integration among these markets has increased over time. According to the author, profitable opportunities from portfolio diversification are therefore limited across major markets, especially during episodes that are marked by a global financial turmoil.

While investigating long-run integration between the United States and many international stock markets, and using the same recursive cointegration technique as in Yunus (2013), Yang *et al.* (2003) found no evidence of cointegration between most of the markets during most of the thirty-two-year sample (1970-2001). The study noted, however, that there is evidence of increasing integration between smaller markets and the United States compared to larger countries. Similarly, Valadkhani and Chancharat (2008) examined the existence of cointegration and causality between stock market price indices and concluded that there is no evidence of a long-run relationship between the stock prices across countries and that potential long-run benefits exist from diversifying investment portfolios internationally to reduce associated systematic risks across countries. Ewing *et al.* (1999), Atteberry and Swanson (1997), and Phengpis and Swanson (2006) found that, although the passage of NAFTA probably enhanced economic interdependence among member countries, they did not detect cointegrating relationship among those markets. The lack of cointegration is also corroborated by Auzairy *et al.* (2012) in their study of the stock market integration of four Asian countries (Malaysia, Thailand, Indonesia, and South Korea) with the world from January 1997 to December 2009. They posited that there is no long-run stock market integration for the four countries and the world market, although there is short-run integration.

Research on financial market integration not only investigates the long-run integration between international equity markets but also questions the stability of that relationship over time. This is important since major events and political changes, including the liberalization of capital controls, deregulations, the growing importance of regional economic integration, and financial crises, may cause structural breaks and, as a result, the dynamic of the integration between equity markets may significantly change over time. Contrary to most previous studies that resorted to the linear modeling technique to investigate financial market integration, some critics argued that equity market integration dynamics should be nonlinear and asymmetric. Market liberalization and financial crises may have induced some persistence, asymmetry, irregularity, and nonlinearity in the stock markets integration process (Jawadi and Arouri 2008; Arouri and Jawadi, 2010; Bekaert, and Harvey, 1995; Li, 2002). The time-varying nature of the long-run integration of equity markets was addressed by Aggarwal and Kyaw (2005), Darrat and Zhong (2005), and Gilmore and McManus (2004) in their investigation of the effect of the passage to NAFTA. They found no evidence of cointegration in the pre-NAFTA period but reported strong evidence in the post-NAFTA period. Bekaert and Harvey (1995), based on a measure of capital market integration arising from a conditional regime-switching model, found that several of emerging markets exhibit time-varying integration. They concluded that while there is a perception that world capital markets have become more integrated, their country-specific investigation suggests that this is not always the case. This is in line with Lopez-Herrera *et al.* (2008) who provided evidence that integration among NAFTA capital markets changes over time, sometimes even decreasing in intensity. Gupta and Donleavy (2009) also demonstrated that correlations are changing over time but noted that correlation may be increasing.

Another line of research investigates the impact of regional trade blocks or economic and monetary unions on financial market integrations. The proliferation of these trade and currency unions is expected to increase the integration of financial markets and to lead to co-movements of the member countries' financial markets. Different views are expressed in the literature on this subject as well, different. In the case of the North American Free Trade Agreement (NAFTA), Phengpis and Swanson (2006) found evidence of a cointegration relation which is time-varying and statistically unstable while using rolling cointegration tests to investigate the evolving nature of the member countries' stock market interdependencies. Ciner (2006) posited that although recent research claimed that NATFA was instrumental in promoting convergence of national markets, the documented cointegration property among those equity markets was in fact confined to a sub-period in the late 1990s. According to the author, the co-movement was caused by the global boom in information technology shares. Based on the Johansen's cointegration method, a stable relation was uncovered between January 1994 and March 2000 but completely disappeared after the March 2000 market crash. Ewing *et al.* (1999) found no evidence of cointegration over the post-USA stock market crash period (1987-1997), even after taking into account the passage of NAFTA. This finding led the authors to conclude that the stock markets of North America were segmented and that the passage of NAFTA has not resulted in greater integration of these stock markets. In a related study, Ewing *et al.* (2001) examined the transmission of the stock return volatility across North American markets during the pre- and post-NAFTA periods using autoregressive conditional heteroskedasticity (ARCH) and vector autoregression (VAR) models and provided evidence on the extent to which cross-market relationships existed in the pre- and post-NAFTA periods. Other studies about NAFTA painted a clearer and more positive picture of the impact of the formation of the regional bloc on its national equity markets. Aggarwal and Kyaw (2005), Darrat and Zhong (2005), and Gilmore and McManus (2004) found strong convergence and cointegration among the equity markets as a result of NAFTA.

There are also studies in the literature also that assessed the degree of integration of financial markets within trade or economic groups in Europe, Asia, Latin America, and other parts of the world. Phengpis *et al.* (2004) investigated the effects of economic convergence among major European Economic and Monetary Union (EMU) member countries on stock market returns in each respective nation. Kim *et al.* (2005) examined the influence of the EMU on the dynamic process of stock market integration using a bivariate exponential general autoregressive conditional heteroskedastic (EGARCH) framework with time-varying conditional correlations. Both studies concluded that economic convergence appears to be an important contributing factor to returns from stock markets in the included EMU countries except Germany and that the increase in stock market integration in Europe over the period 1999-2003 has been significantly driven, in part, by macroeconomic convergence associated with EMU. Cointegration of the financial market in the European Union (EU) and EMU areas are also supported by other studies, including Serletis and King (1997), Fratzscher (2002), Hardouvelis *et al.* (2006), and Croci (2004). In a related study, Laopodis (2005) investigated the possibility of cointegration between the United States and eleven European equity markets before and after the convergence period of 1995. The author found at least one cointegrating vector that emerged in either period for the European Union markets, but no cointegration among them surfaced during the Euro introduction period of 1999. The finding of no strong cointegration by Laopodis (2005) is echoed few years later by da Fonseca (2008). The latter shows that the introduction of a single currency did not affect the nature of the long-term relationships between the variables, and that both European and non-European international factors are necessary to explain the international integration of the national stock markets under analysis.

Calvi (2010) assessed and compared the degree of integration of financial markets within Europe (seven countries) and East Asia (eleven countries). The analyses, based on cointegration and Granger causality techniques, confirmed that financial integration is significantly more advanced in Europe compared to East Asia. Gharleghi *et al.* (2015) investigated financial integration in the Association of Southeast Asian Nations (ASEAN) using data from 1985 to 2010 and also divided the sample into two groups of pre-crisis (1985-1996) and post-crisis (1999-2010). The empirical analysis revealed that there was a unique cointegrating vector in the post-crisis period only, suggesting a long-run equilibrium relationship between ASEAN exchange rates. In studying the impact of regional integration in the Caribbean, Lorde *et al.* (2009) and Harrison and Moore (2010) investigated the co-movements among selected Caribbean stock markets. The two studies failed to find any evidence of co-movements among the stock markets investigated.

Researchers continued to investigate the degree of integration of financial markets in different parts of the world by using recent data. Using structural VAR analysis, Shu *et al.* (2018) compared spillovers from the USA and Chinese financial markets to markets in Asia and the Pacific. Their results revealed that, in normal times, China's influence in the equity market has risen to a level close to that of the United States, although the relative impact of the United States became stronger in crisis periods. While analyzing the long-run relationships amongst six Latin American stock markets, Vides (2021) applied fractional cointegration to recent observations for the period 2002 to 2019. The results showed there are four cointegrating vectors among the six equity markets, suggesting that Latin American stock markets are not fully nor perfectly integrated. In addition, based on recent data and using fractional integration and fractional cointegration methods, Caporale *et al.* (2021) examined stock market integration between the five ASEAN countries and both the USA and China in turn. Their main findings showed that all stock indices exhibit long-range dependence, and there is cointegration between the five ASEAN countries and the USA but almost none between the former and China, except between Indonesia and China in the case of the financial sector.

## 3. Methodology and model

### 3.1. Model specification, stationarity, and cointegration Analysis

In this section, we investigate the co-movements among the aggregate stock indices of countries within different regions: NAFTA countries, the Eurozone countries, and a selected group of OECD countries. Following Toulaboue (2016), we estimate the following model as shown in Equation 1 for each country group to investigate the long-run cointegration relationships among the stock market indices.

$$IndCnt_{1t} = \lambda + \varphi_2 IndCnt_{2t} + \varphi_3 IndCnt_{3t} + \varphi_4 IndCnt_{4t} + \cdots + \upsilon_t \quad (1)$$

The variable *IndCnt*$_i$ is the stock market index of country *i* (*i* = 1, 2, 3, …), and $\varphi_j$ (j = 2, 3, 4 …) are the long-run coefficients. $\lambda$ is the constant term, whereas $\upsilon_t$ is the white noise error term.

We rely on both the traditional Johansen (1988) cointegration approach and the more advanced autoregressive distributed lag model (ARDL) bounds testing approach (developed by Pesaran and Shin (1999) and Pesaran *et al.* (2001) to investigate the long-run cointegration relations. The Johansen test is widely used to examine cointegration. An important precondition that must be satisfied prior to executing the Johansen cointegration test is that all the variables under consideration must have the same order of integration – integrated of order one - I(1). Cointegration is defined as a situation where a linear combination of two or more non-stationary time series is stationary, therefore implying the existence of a long-run equilibrium amongst the variables. In the case of financial markets, the more cointegrating relations are, the higher the cointegration amongst the financial markets under investigation.

Except for its stringent condition for all the variables under investigation to have the same order of integration, the Johansen test is a superior test for cointegration and has all the desirable statistical properties. Johansen proposes two different likelihood ratio tests for scrutinizing the cointegration relationship, namely the trace statistics and the maximum eigenvalue statistics. The null hypothesis of the two tests is that there are at most “r cointegrating relations”. The alternative is that there are “r cointegrating relations” in the case of the trace test and “r+1 cointegrating relations” in the case of the eigenvalue test (Johansen, 1988).

For robustness check and because of its numerous advantages, a second test of cointegration, namely the ARDL bounds testing approach was applied. It has been claimed in the literature that the bounds testing technique has more attractive features compared to the Johansen cointegration approach. The bounds test can be applied to any set of data, whether purely I(0), purely I(1), or a mixture of the two. The bounds test approach to cointegration is, therefore, easier to apply compared to other approaches as its bypasses the use of unit root pre-testing to establish orders of integration of the independent variables and classifies these into I(0) and I(1), (Pesaran *et al.* 2001). This procedure allows different variables to be assigned different lag lengths as they enter the model and is more robust even in small sample cases (Pesaran *et al.* 2001; Narayan, 2005). According to Haug (2002), the ARDL bounds testing approach is more suitable and provides better results for small sample sizes, and the short and long-run parameters are estimated simultaneously.

To implement the bounds testing approach to investigate the long-run relationship amongst the variables, Equation 1 is modeled as a conditional or unrestricted ARDL, and it takes the form as in Equation 2.

$$\begin{aligned}\Delta IndCnt_{1t} = \sigma_0 + \delta_1 IndCnt_{1t-1} + \delta_2 IndCnt_{2t-1} + \delta_3 IndCnt_{3t-1} + \cdots \\ + \textstyle\sum_{j=1}^{m} \theta_1 \Delta IndCnt_{1t-i} + \sum_{j=1}^{m} \theta_2 \Delta IndCnt_{2t-i} \\ + \textstyle\sum_{j=1}^{m} \theta_3 \Delta IndCnt_{3t-i} + \ldots + \xi_t\end{aligned} \quad (2)$$

, where $\sigma_0$ is the drift term, $\delta_j$ (j = 1, 2, 3, …) are the long-run coefficients, $\xi_t$ is the white noise error term, and $m$ is the lag length. The lag length can be determined based on the Akaike Information Criteria (AIC) or Schwarz Information Criteria (SIC) through the process of setting a priori a maximum lag length, with the optimal lag length being the one with the lowest AIC or SIC value.

Now the question is whether we can reject the hypothesis that the long-run coefficients $\delta_j$ (j = 1, 2, 3, …) in the ARDL model (Equation 2) are jointly zero. We resort to the Wald test to investigate this joint hypothesis (Equations 3 and 4) to determine whether the variables in Equation 2 are cointegrated. This involves conducting an F-test on the lagged levels of the independent variables in Equation 2.

$$\delta_1 = \delta_2 = \delta_3 = \cdots = 0 \tag{3}$$
$$\delta_1 \neq \delta_2 \neq \delta_3 \neq \cdots \neq 0 \tag{4}$$

The calculated Wald F-statistic is then compared to the critical values (tabulated by Pesaran *et al.* (2001)). If the calculated F-statistic falls above the upper critical value, this establishes evidence of cointegration amongst the variables of interest. If it falls below the lower bound critical value, no cointegration is detected amongst the variables. If the F-statistic falls between the lower bound and upper bound critical values, then the cointegration test is inconclusive (Pesaran *et al.* 2001).

Assuming that, as a result of the cointegration tests, a long-run relationship is detected amongst the variables, the next step is to estimate the dynamic short-run relationship amongst the variables. Causal relations among variables can be examined within the framework of VECM with cointegrated variables as in Equation 5.

$$\Delta IndCnt_{1t} = c_0 + \psi ECT_{t-1} + \sum_{j=1}^{p} \tau_1 \Delta IndCnt_{1t-i} + \sum_{j=1}^{p} \tau_2 \Delta IndCnt_{2t-i} + \sum_{j=1}^{p} \tau_3 \Delta IndCnt_{3t-i} + \cdots + \eta_t \tag{5}$$

The variable $ECT_{t-1}$ is the error correction term and its coefficient ($\psi$) is a measure of the speed of adjustment to the long-run equilibrium. A negative and statistically significant $ECT_{t-1}$ is a sign that a long-run causality is detected. The individual coefficients of the lagged terms $\tau_j$ (j = 1, 2, 3, …) are the short-run dynamic coefficients. Their statistical significance depicts short-run causality. The coefficients $c_0$ and $\eta_t$ are the constant term and the white noise error term, respectively.

Finally, a Granger causality test is performed using the Wald test under the VECM environment to investigate causal relationships amongst the aggregate indices. Granger (1969) defined a concept of causality which has become popular in recent years in investigating causal relationships among variables. In the case of two time-series (X and Y), X is said to Granger-cause Y if Y can be better predicted using the histories of both X and Y than it can by using the history of Y alone.

### 3.2. Methodology and data

As stated in the previous section, the main objective of this study is to examine and compare the linkages of stock indices in different country groups: NAFTA, the Eurozone, and selected OECD countries. To analyze the international integration of the national stock markets under investigation, different methodologies, and steps are applied. First, we provide summary statistical characteristics of the national stock market prices in the different market groups, followed by tests of stationarity of all the time series included in the models using the Augmented Dickey-Fuller (ADF) test. The results of these unit root tests to check the order of integration of the stock price indices will inform us later about the use of the Johansen cointegration test. Simple correlation matrix analyses are then conducted to determine the correlation amongst the national stock price indices in the different markets. Using two different methods of cointegration analysis (Johansen cointegration approach and the autoregressive distributed lag bounds testing approach), we then investigate long-term relationships among the national stock indices of the three country groups to assess the level of their integration. Finally, the vector-error correction model (VECM) and Granger causality test are used to garner additional useful information. The VECM is applied to investigate the short-run (and long-run) dynamics of the stock market indices, while the Granger causality test is applied to investigate the existence and direction of causality between the market indices.

Different national stock market indices are used in this study. The data consist of monthly observations of the national stock market indices from NAFTA, the Eurozone[2], and a selected group of OECD countries. Instead of daily or weekly (high frequency) data, monthly data series are used in this study in order to appropriately assess the existence of cointegration (Bailey and Stulz, 1990). All the indices used are converted into natural logarithms to allow for better comparison among the countries. They are all obtained from OECD (2019) (share prices indicator) and cover the period of 1994-2018 for NAFTA, 2000-2018 for the Eurozone, and 2000-2018 for the selected OECD countries.[3] Each national stock market is represented by its national stock index.

## 4. Empirical results and analysis

This section presents the empirical results of this study. We first explore the descriptive statistics and correlation matrices.

### 4.1. Descriptive statistics and correlation matrix

Table 1 shows the descriptive statistics for the national stock market indices for different country groups. The table shows that within the NAFTA region, the Mexican market exhibits a lower mean (index levels) but higher volatility and risk compared to the Canadian and USA markets. The higher risk associated with the Mexican stock market is compensated for by a higher mean return (first difference of the log index level – not shown) compared to the Canadian and USA markets.

Taking the return series as a benchmark, the mean return for the Mexican stock market is roughly twice as high as the mean returns for the Canadian and USA stock markets. Regarding skewness, all three stock market series are negatively skewed. For the Eurozone, the index level mean is around 2.00 across the eleven countries, with volatility highest in Austria, Ireland, and Luxemburg, and lowest in France, Portugal, and Spain. Finland, Luxembourg, and Italy experienced the highest mean returns while Austria, Belgium, and Ireland trail with the lowest.

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The series in the Eurozone are mostly negatively skewed. In the group of selected OECD countries, the means of the index levels range from 1.647 (in Turkey) to 2.093 (in Israel), with the highest volatility detected in Turkey, Israel, and Chile, and the lowest in Australia, Japan, and New Zealand. The highest mean returns are uncovered for the Turkish, Chilean, and Koran markets, while Israel, Japan, and New Zealand represent the tail dogs for that group. Except for Israel and New Zealand, the index series are negatively skewed.

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[2] Countries included in the Eurozone sample in this study are members of the Eurozone since 1999.

[3] Share price indices are calculated from the prices of common shares of companies traded on national or foreign stock exchanges. They are usually determined by the stock exchange, using the closing daily values for the monthly data, and normally expressed as simple arithmetic averages of the daily data.

**Table 1. Descriptive statistics**

**NAFTA**

| | USA | Canada | Mexico |
|---|---|---|---|
| Mean | 1.812 | 1.825 | 1.511 |
| Minimum | 1.390 | 1.447 | 0.569 |
| Maximum | 2.094 | 2.059 | 2.067 |
| Median | 1.820 | 1.870 | 1.644 |
| Range | 0.704 | 0.611 | 1.497 |
| Std Error | 0.010 | 0.010 | 0.026 |
| Std Dev. | 0.170 | 0.170 | 0.455 |
| Variance | 0.029 | 0.028 | 0.207 |
| Skewness | -0.708 | -0.562 | -0.335 |
| Kurtosis | 0.095 | -0.813 | -1.390 |
| Count | 300 | 300 | 300 |

**Eurozone**

| | AUT | BEL | DEU | ESP | FIN | FRA | IRL | ITA | LUX | NLD | PRT |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Mean | 1.977 | 1.862 | 1.851 | 1.961 | 1.958 | 1.919 | 1.902 | 2.013 | 2.054 | 1.937 | 2.001 |
| Minimum | 1.672 | 1.571 | 1.500 | 1.752 | 1.726 | 1.686 | 1.523 | 1.774 | 1.769 | 1.669 | 1.767 |
| Maximum | 2.294 | 2.052 | 2.072 | 2.196 | 2.304 | 2.068 | 2.195 | 2.242 | 2.399 | 2.131 | 2.240 |
| Median | 1.997 | 1.843 | 1.860 | 1.962 | 1.951 | 1.928 | 1.932 | 1.998 | 2.015 | 1.944 | 2.002 |
| Range | 0.622 | 0.482 | 0.572 | 0.444 | 0.578 | 0.382 | 0.672 | 0.468 | 0.630 | 0.462 | 0.473 |
| Std Error | 0.011 | 0.008 | 0.008 | 0.006 | 0.008 | 0.006 | 0.010 | 0.008 | 0.009 | 0.007 | 0.006 |
| Std Dev. | 0.166 | 0.117 | 0.128 | 0.095 | 0.117 | 0.094 | 0.154 | 0.116 | 0.137 | 0.102 | 0.093 |
| Variance | 0.028 | 0.014 | 0.016 | 0.009 | 0.014 | 0.009 | 0.024 | 0.013 | 0.019 | 0.010 | 0.009 |
| Skewness | -0.235 | -0.206 | -0.391 | 0.364 | 0.597 | -0.286 | -0.423 | 0.245 | 0.687 | -0.123 | -0.065 |
| Kurtosis | -0.758 | -1.023 | -0.518 | 0.187 | 0.156 | -0.892 | -0.720 | -0.724 | -0.034 | -0.777 | 0.170 |
| Count | 228 | 228 | 228 | 228 | 228 | 228 | 228 | 228 | 228 | 228 | 228 |

**Selected OECD Countries**

| | AUS | CHL | CZE | ISL | JPN | KOR | NZL | SWE | TUR |
|---|---|---|---|---|---|---|---|---|---|
| Mean | 1.903 | 1.813 | 1.950 | 2.093 | 1.889 | 1.827 | 1.909 | 1.793 | 1.647 |
| Minimum | 1.698 | 1.266 | 1.528 | 1.534 | 1.673 | 1.399 | 1.751 | 1.436 | 0.506 |
| Maximum | 2.085 | 2.186 | 2.276 | 2.844 | 2.082 | 2.100 | 2.153 | 2.068 | 2.160 |
| Median | 1.926 | 1.885 | 1.991 | 2.049 | 1.900 | 1.914 | 1.891 | 1.798 | 1.755 |
| Range | 0.387 | 0.920 | 0.748 | 1.310 | 0.409 | 0.701 | 0.402 | 0.633 | 1.655 |
| Std Error | 0.007 | 0.017 | 0.012 | 0.021 | 0.007 | 0.013 | 0.007 | 0.010 | 0.024 |
| Std Dev. | 0.106 | 0.259 | 0.181 | 0.326 | 0.115 | 0.203 | 0.104 | 0.160 | 0.377 |
| Variance | 0.011 | 0.067 | 0.033 | 0.107 | 0.013 | 0.041 | 0.011 | 0.026 | 0.142 |
| Skewness | -0.309 | -0.458 | -0.522 | 0.661 | -0.246 | -0.596 | 0.565 | -0.146 | -0.733 |
| Kurtosis | -1.233 | -1.225 | -0.534 | -0.408 | -1.162 | -1.076 | -0.784 | -0.756 | -0.491 |
| Count | 240 | 240 | 240 | 240 | 240 | 240 | 240 | 240 | 240 |

We now turn to the simple correlation coefficients to examine the level of association between the stock indices in different country groups. Table 2 presents the results. Starting with NAFTA, the correlations among the indices of the three markets are remarkably high, ranging from 0.890 to 0.961. This high correlation amongst the NAFTA stock indices is not surprising and may be due to a few factors.

NAFTA is a free trade area that allows member countries to trade with each other under limited restrictions. Additionally, neighboring countries tend to exhibit higher pairwise correlations. The close proximity and the establishment of the free trade area result in a high degree of economic interdependence and contribute to increased capital flows within the region. Economic theory and empirical evidence show that enlarged and regional markets create new opportunities for market participants, thereby contributing to optimal use of financial resources as investors can easily switch investments across markets and increase global portfolio diversification. The characteristics and nature of NAFTA are therefore expected to generate increased financial integration, convergence, and higher correlations amongst the related financial markets. It is important, however, to point out that the degree of risk reduction, resulting from a global portfolio diversification in a region with a high correlation of returns, will be severely diminished. The significantly high correlations of NAFTA markets imply that expected portfolio diversification benefits of investing in NAFTA markets are seriously hampered, and that global investors should be able to reduce portfolio risk more if they diversify internationally outside NAFTA rather than within.

**Table 2. Correlation matrix**

**NAFTA**

| | USA | Canada | Mexico |
|---|---|---|---|
| USA | 1.000 | | |
| Canada | 0.958 | 1.000 | |
| Mexico | 0.890 | 0.961 | 1.000 |

**Eurozone**

| | AUT | BEL | DEU | ESP | FIN | FRA | IRL | ITA | LUX | NLD | PRT |
|---|---|---|---|---|---|---|---|---|---|---|---|
| AUT | 1.000 | | | | | | | | | | |
| BEL | 0.788 | 1.000 | | | | | | | | | |
| DEU | 0.493 | 0.852 | 1.000 | | | | | | | | |
| ESP | 0.751 | 0.673 | 0.482 | 1.000 | | | | | | | |
| FIN | 0.245 | 0.576 | 0.706 | 0.576 | 1.000 | | | | | | |
| FRA | 0.502 | 0.851 | 0.905 | 0.629 | 0.881 | 1.000 | | | | | |
| IRL | 0.399 | 0.673 | 0.470 | 0.443 | 0.646 | 0.730 | 1.000 | | | | |
| ITA | 0.161 | 0.298 | 0.198 | 0.572 | 0.742 | 0.580 | 0.737 | 1.000 | | | |
| LUX | 0.472 | 0.409 | 0.364 | 0.844 | 0.657 | 0.543 | 0.312 | 0.660 | 1.000 | | |
| NLD | 0.243 | 0.670 | 0.796 | 0.469 | 0.925 | 0.938 | 0.732 | 0.674 | 0.525 | 1.000 | |
| PRT | 0.756 | 0.754 | 0.707 | 0.889 | 0.685 | 0.774 | 0.424 | 0.467 | 0.793 | 0.613 | 1.000 |

**Selected OECD Countries**

| | AUS | CHL | CZE | ISL | JPN | KOR | NZL | SWE | TUR |
|---|---|---|---|---|---|---|---|---|---|
| AUS | 1.000 | | | | | | | | |
| CHL | 0.849 | 1.000 | | | | | | | |
| CZE | 0.830 | 0.706 | 1.000 | | | | | | |
| ISL | 0.166 | -0.277 | 0.340 | 1.000 | | | | | |
| JPN | 0.425 | 0.032 | 0.229 | 0.571 | 1.000 | | | | |
| KOR | 0.881 | 0.959 | 0.742 | -0.186 | 0.162 | 1.000 | | | |
| NZL | 0.790 | 0.591 | 0.519 | 0.337 | 0.687 | 0.627 | 1.000 | | |
| SWE | 0.830 | 0.761 | 0.520 | -0.060 | 0.569 | 0.809 | 0.772 | 1.000 | |
| TUR | 0.897 | 0.966 | 0.749 | -0.150 | 0.158 | 0.947 | 0.659 | 0.813 | 1.000 |

The characteristics of NAFTA are, to a certain degree, similar to those of the Eurozone. The Eurozone, by providing the same currency to member countries, facilitates trade and financial transactions. The increased level of transactions within the zone, coupled with the proximity argument outlined above, should contribute to the increased integration of markets in that geographic economic region. The correlation coefficients of markets in the Eurozone largely confirm that claim. The coefficients range from 0.161 to 0.938, with about 70 percent of the coefficients exceeding 0.500. The country pairs with the highest market correlations are France/Netherlands, Finland/Netherlands, and France/Germany. Country pairs that exhibit the lowest correlations are Austria/Italy, Austria/Netherlands, and Italy/Germany.

The selected OECD countries as a group, contrary to NAFTA and the Eurozone countries in our analysis do not, belong to a common regional economic integration or currency union. This may lead one to hypothesize that markets in those countries may not experience an intensification of cross-border movement of capital (and thereby a high level of correlation of their markets) at the level evidenced in the case of NAFTA and the Eurozone. Table 2 shows that the correlation coefficients of markets in the selected OECD countries are relatively weaker and, in some cases, negative. The coefficients range from -0.277 to 0.966. The Israeli market experienced the weakest correlation with other countries in that group.

Although we detected evidence of correlation among all national stock indices in this study, it should be concluded that the correlation coefficients are much stronger in areas that belong to regional economic integration (NAFTA and Eurozone). It should also be reminded that the main objective of this study is an investigation of market integration, and the information provided by correlation coefficients is not enough to claim the existence of market integration

(Kasa, 1992). A high correlation between two variables does not guarantee cointegration, nor does cointegration imply a high correlation. In the next section, we explore two, more systematic ways to investigate market integration.

### 4.2. Cointegration analysis

In this section, we investigate long-term relationships among the national stock indices of the three country groups to assess the level of their integration. The existence of cointegration is an indication that a long-run relationship exists amongst the variables of interest. We rely on two different methods, the Johansen cointegration approach and the autoregressive distributed lag (ARDL) bounds testing approach, to test for cointegration of the market indices.

As stated, an important condition for applying the Johansen cointegration test is that all variables in the system should be nonstationary, integrated of the same order. The stationarity tests of all the time series included in the models are conducted using the Augmented Dickey-Fuller (ADF) unit root test to determine their order of integration. The selection of the optimal lags in the test specification for each variable is obtained through a search process (i.e., the minimum Schwarz Information Criteria, SIC). Overall, the results of the unit root tests (not reported due to the size of the data) exhibit evidence that all the variables in the models are nonstationary in levels but stationary in first differences, an indication that they are integrated of the same order one, I(1). This provides the basis for cointegration analysis.

We move on to apply the Johansen cointegration test to Equation 1 for each country group. Again, the lags are selected based on the Schwartz Information Criteria (SIC). For each country group, the test is conducted for the full sample and subsamples. Table 3 reports the results of the Johansen cointegration tests. For NAFTA, considering the results of the full sample (1994-2018), the trace and maximum-eigenvalue statistics clearly show that the null hypothesis of no cointegration cannot be rejected at the 5% significance level. The results, therefore, show no evidence of cointegration amongst the NAFTA markets. These results are consistent with Ewing *et al.* (1999), Atteberry and Swanson (1997), and Phengpis and Swanson (2006), who found no cointegrating relationship among the NAFTA markets.

The results are also in contrast with studies by Aggarwal and Kyaw (2005), Darrat and Zhong (2005), and Gilmore and McManus (2004)), which reported convergence and cointegration among the equity markets as a result of NAFTA. Based on the two polar views and the emerging theory that market cointegration relations are time-varying (Lopez-Herrera *et al.* 2008; Gupta and Donleavy, 2009), we apply the cointegration analysis to three sub-periods: 1994-1999, 2001-2007, and 2008-2018 to account for possible structural breaks in the series. The results (Table 3) are telling. In each of the three sub-periods, there is clear evidence of cointegration relations amongst the markets at the 5% significance level based on the trace and maximum-eigenvalue statistics.

Given the no cointegration results of the full sample and the resounding cointegration relations unveiled in the sub-periods, one can suspect the existence of structural breaks in the data series. The cointegration nature of these markets as revealed by the results in this study, supports previous claims that NAFTA was instrumental in promoting convergence of those national markets.

The Johansen cointegration tests for the Eurozone and the selected OECD countries are more consistent across samples. Table 3 shows that, for both country groups, the trace and maximum-eigenvalue statistics reject the null hypothesis of no cointegration at the 5% significance level for the full sample as well as the sub-periods.

These results therefore show evidence of cointegration and confirm the existence of long run relations and interdependence amongst the stock markets of the Euro Zone and the selected OECD countries. The results largely align with Serletis and King (1997), Fratzscher (2002), Hardouvelis *et al.* (2006), and Croci (2004) who supported cointegration of financial market in the European Union (EU) and EMU.

**Table 3. Johansen cointegration tests**

| Ho: r | Trace Statistic | Critical Value 5% | Max-Eigenv. Statistic | Critical Value 5% |
|---|---|---|---|---|
| **NAFTA** | | | | |
| *Full Sample: 1994-2018* | | | | |
| None | 27.1120 | 29.7971 | 18.9717 | 21.1316 |
| At most 1 | 8.1403 | 15.4947 | 6.9053 | 14.2646 |
| At most 2 | 1.2350 | 3.8415 | 1.2350 | 3.8415 |
| *Sub-Sample: 1994-1999* | | | | |
| None | 30.2260* | 29.7971 | 24.8765* | 21.1316 |
| At most 1 | 5.3495 | 15.4947 | 5.1152 | 14.2646 |
| At most 2 | 0.2344 | 3.8415 | 0.2344 | 3.8415 |
| *Sub-Sample: 2001-2007* | | | | |
| None | 34.0333* | 29.7971 | 25.1139* | 21.1316 |
| At most 1 | 8.9194 | 15.4947 | 8.6951 | 14.2646 |
| At most 2 | 0.2243 | 3.8415 | 0.2243 | 3.8415 |
| *Sub-Sample: 2008-2018* | | | | |
| None | 47.9365* | 29.7971 | 26.1127* | 21.1316 |
| At most 1 | 21.8238* | 15.4947 | 14.9398* | 14.2646 |
| At most 2 | 6.8840* | 3.8415 | 6.8840* | 3.8415 |
| **EURO Zone** | | | | |
| **Ho: r** | **Trace Statistic** | **Critical Value 5%** | **Max-Eigenv. Statistic** | **Critical Value 5%** |
| *Full Sample: 2000-2018* | | | | |
| None | 324.2522* | 285.1425 | 74.1447* | 70.5351 |
| At most 1 | 250.1075* | 239.2354 | 58.4710 | 64.5047 |
| At most 2 | 191.6365 | 197.3709 | 49.1009 | 58.4335 |
| *Sub-Sample: 2000-2007* | | | | |
| None | 397.1362* | 285.1425 | 99.5085* | 70.5351 |
| At most 1 | 297.6277* | 239.2354 | 84.7483* | 64.5047 |
| At most 2 | 212.8794* | 197.3709 | 61.2815* | 58.4335 |
| *Sub-Sample: 2008-2018* | | | | |
| None | 415.7194* | 285.1425 | 94.4110* | 70.5351 |
| At most 1 | 321.3083* | 239.2354 | 66.7612* | 64.5047 |
| At most 2 | 254.5472* | 197.3709 | 64.3626* | 58.4335 |
| **Selected OECD Countries** | | | | |
| **Ho: r** | **Trace Statistic** | **Critical Value 5%** | **Max-Eigenv. Statistic** | **Critical Value 5%** |
| *Full Sample: 2000-2018* | | | | |
| None | 290.4777* | 197.3709 | 83.0809* | 58.4335 |
| At most 1 | 207.3968* | 159.5297 | 54.5830* | 52.3626 |
| At most 2 | 152.8139* | 125.6154 | 45.6860 | 46.2314 |
| *Sub-Sample: 2000-2007* | | | | |
| None | 349.6774* | 197.3709 | 92.9003* | 58.4335 |
| A most 1 | 256.7771* | 159.5297 | 77.4435* | 52.3626 |
| At most 2 | 179.3336* | 125.6154 | 65.7768* | 46.2314 |
| *Sub-Sample: 2008-2018* | | | | |
| None | 255.7404* | 197.3709 | 76.6163* | 58.4335 |
| At most 1 | 179.1240* | 159.5297 | 51.3359 | 52.3626 |
| At most 2 | 127.7882* | 125.6154 | 37.2601 | 46.2314 |

**Note:** * denotes rejection of the null hypothesis at the 5% level.

To investigate the robustness of the cointegration analyses of the financial markets, we use the ARDL bounds testing approach. The results are reported in Table 4. Table 5 compares the Johansen cointegration results with those of the ARDL bounds testing approach.

**Table 4. Bounds test for cointegration: F-statistics and critical values**

**NAFTA**

| **F- Statistic** | | | | **Critical values** | | |
|---|---|---|---|---|---|---|
| *Full Sample 1994-2018* | *Sub-Sample 1994-1999* | *Sub-Sample 2000-2007* | *Sub-Sample 2008-2018* | Signif. Level | *Lower Bound* I(0) | *Upper Bound* I(1) |
| 1.5966 | 7.2662* | 5.2280* | 4.1060** | 10% | 2.63 | 3.35 |
| | | | | 5% | 3.10 | 3.87 |
| | | | | 1% | 4.13 | 5.00 |

**EURO Zone**

| **F- Statistic** | | | **Critical values** | | |
|---|---|---|---|---|---|
| *Full Sample 2000-2018* | *Sub-Sample 2000-07* | *Sub-Sample 2009-2018* | Signif. Level | *Lower Bound* I(0) | *Upper Bound* I(1) |
| 1.9954 | 6.3996* | 3.1010** | 10% | 1.76 | 2.77 |
| | | | 5% | 1.98 | 3.04 |
| | | | 1% | 2.41 | 3.61 |

**Selected OECD Countries**

| **F- Statistic** | | | **Critical values** | | |
|---|---|---|---|---|---|
| *Full Sample 2000-2018* | *Sub-Sample 2000-2007* | *Sub-Sample 2008-2018* | Signif. Level | *Lower Bound* I(0) | *Upper Bound* I(1) |
| 2.4097 | 5.6001* | 3.1903** | 10% | 1.85 | 2.85 |
| | | | 5% | 2.11 | 3.15 |
| | | | 1% | 2.62 | 3.17 |

**Note:** ***, **, and * denotes the significance levels at 1%, 5%, and 10% respectively.

Considering the full sample (in Table 4), the ARDL bounds test calculated that F-statistics are lower than the upper bound of the critical values at the 1%, 5%, and 10% levels for each of the three country groups (NAFTA, Eurozone, and the selected OECD countries). This is evidence of no cointegration and shows that no long-run relationship exists amongst the markets in each of those three country groups based on the full sample. Considering the NAFTA case, these bounds test results are consistent with those based on the Johansen tests reported in Table 3 for the full sample. However, the bounds test results for the other two country groups (Eurozone and selected OECD countries) contradict those obtained by the Johansen test. While the Johansen test shows evidence of cointegration, the ARDL bound testing approach rejects such evidence for the two country groups. Considering the full sample, the results therefore unequivocally point to no cointegration in the case of NAFTA, while they are mixed in the case of the other two country groups.

Considering the sub-periods, the ARDL bound test calculated that F-statistics are greater than the upper bound of the critical values at the 1%, 5%, and 10% levels for each of the three country groups (NAFTA, Eurozone, and selected OECD countries). This is an indication of cointegration. Clear and consistent results, therefore, emerged when sub-periods are considered. The results from both the Johansen and the ARDL bound testing approaches unveil evidence of cointegration and long-run relations in each of the sub-periods for the three country groups.

It can be concluded that financial markets in the OECD countries, especially the three country groups considered in this study, are cointegrated and that those markets follow the same long-run stochastic paths. The results also lend support to the dynamic nature of the long-run integration of equity markets and those markets may exhibit time-varying long-run interdependence.

This cointegration of markets, of course, has important regional and international implications for global portfolio investors, as gains from portfolio diversification are limited or may not materialize across these markets.

**Table 5. Comparison of Johansen and ARDL bound test cointegration results**

| | **Full Sample** | | **Sub-Periods** | | | | | |
|---|---|---|---|---|---|---|---|---|
| **NAFTA** | 1994-2018 | | 1994-1999 | | 2001-2007 | | 2008-2018 | |
| | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** |
| | No | No | Yes | Yes | Yes | Yes | Yes | Yes |
| **Eurozone** | 2000-2018 | | 1994-1999 | | 2001-2007 | | 2000-2018 | |
| | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** |
| | Yes | No | --- | --- | Yes | Yes | Yes | Yes |
| **Other OECD** | 2000-2018 | | 1994-1999 | | 2001-2007 | | 2008-2018 | |
| | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** | **JC** | **BT** |
| | Yes | No | --- | --- | Yes | Yes | Yes | Yes |

**Notes:** JC denotes Johansen cointegration, whereas BT is the ARDL bounds test cointegration.

### 4.3. Vector-error correction model and Granger causality

After investigating the cointegration and long-run relationships amongst the groups of markets of interest, it will be insightful to also examine the short-run dynamic interactions amongst them to help provide additional useful information. This is done in this study using the VECM and the Granger causality test. According to the Granger Representation Theorem, if two variables are cointegrated, then the relationship between the two can be expressed as a vector-error correction model. The VECM analysis is conducted for the NAFTA markets only in this study and for the most recent sub-period, 2008-2018. The choice of optimal lags is selected based on the AIC, and two lags are chosen. The results are presented in Table 6.

**Table 6. NAFTA - VECM estimations (sample: 2008-2018)**

| **Repressors** | **ΔIndUSA** | **ΔIndCAN** | **ΔIndMEX** |
|---|---|---|---|
| Constant | 0.0005<br>(0.430) | 0.0001<br>(0.071) | 0.0011<br>(0.802) |
| ΔIndUSA(-1) | -0.216<br>(-1.151) | -0.2590<br>(-1.488) | -0.5552$^{*}$<br>(-2.798) |
| ΔIndUSA(-2) | -0.5117$^{*}$<br>(-2.630) | -0.4473$^{**}$<br>(-2.478) | -0.3445$^{***}$<br>(-1.674) |
| ΔIndCAN(-1) | 0.3684$^{**}$<br>(2.029) | 0.3476$^{**}$<br>(2.063) | 0.5049*<br>(2.630) |
| ΔIndCAN(-2) | 0.4829$^{*}$<br>(2.610) | 0.5070$^{*}$<br>(2.953) | 0.3254$^{***}$<br>(1.663) |
| ΔIndMEX(-1) | 0.2239$^{* * *}$<br>(1.809) | 0.2291$^{* *}$<br>(1.995) | 0.3468$^{*}$<br>(2.650) |
| ΔIndMEX(-2) | -0.0332<br>(-0.266) | 0.0551<br>(0.4759) | -0.0854<br>(-0.646) |
| ECM(-1) | -0.1197$^{*}$<br>(-4.278) | -0.0616$^{**}$<br>(-2.372) | -0.0663$^{**}$<br>(-2.242) |
| $R^2$ | 0.31 | 0.26 | 0.19 |
| $R^2$ Adj. | 0.27 | 0.21 | 0.15 |
| LM(1) | 0.455<br>(0.5000) | 0.9623<br>(0.3266) | 0.2906<br>(0.5898) |
| LM(2) | 0.831<br>(0.660) | 5.0769<br>(0.0790) | 3.0157<br>(0.2214) |
| ARCH | 0.046<br>(0.977) | 0.211<br>(0.900) | 2.148<br>(0.3417) |

**Notes:** IndUSA is the USA index, IndCAN is the Canadian Index, and IndMEX is the Mexican Index. *, **, and *** indicates the significance levels at 10%, 5%, and 1% respectively.

The reliability of these error correction models is documented through a number of diagnostic tests, including the Breusch-Godfrey LM test for serial correlation, CUSUM tests for stability, and the ARCH test for heteroscedasticity. The results of these diagnostic tests, also presented in Table 6 and Figure 1 for the case of Ramsey RESET test, show that the short-run dynamic models are well specified. This suggests that the models do not suffer from serial correlation and heteroscedasticity, and that they are stable at reasonable significance levels. Cointegration has been established amongst the NAFTA markets for the sample of interest (2008-2018), implying that a long-run equilibrium relation exists amongst those markets. Therefore, the estimated coefficients of the error-correction term, ECM(-1), from the VECM models should be negative and significant. Table 6 shows this to be true. The coefficients are negative across all models and statistically significant. These negative signs and statistical significance are further evidence of cointegration, and there exists some adjustment mechanism that prevents the variables from drifting apart in the long run, ensuring a long-run relationship among them. These results boldly addressed and confirmed the results of the main (cointegration) query in this study. The magnitude of the coefficients of the error correction term, a measure of the speed of adjustment of short-run fluctuations toward long-run equilibrium, is 0.1197 for the USA, 0.0616 for Canada, and 0.0663 for Mexico. All three countries appear to adjust to correct for past systematic disequilibrium.

In the short run, the USA equation suggests that one lagged IndCAN and IndMEX and two lagged IndUSA and IndCAN variables have significant impact on USA market index. For the Canadian equation, except for one lagged IndUSA and two lagged IndMEX, the other variables have significant short-run effect on the Canadian stock markets. For the Mexican equation, except for two lagged IndMEX, short run impacts are detected for all variables in the equation.

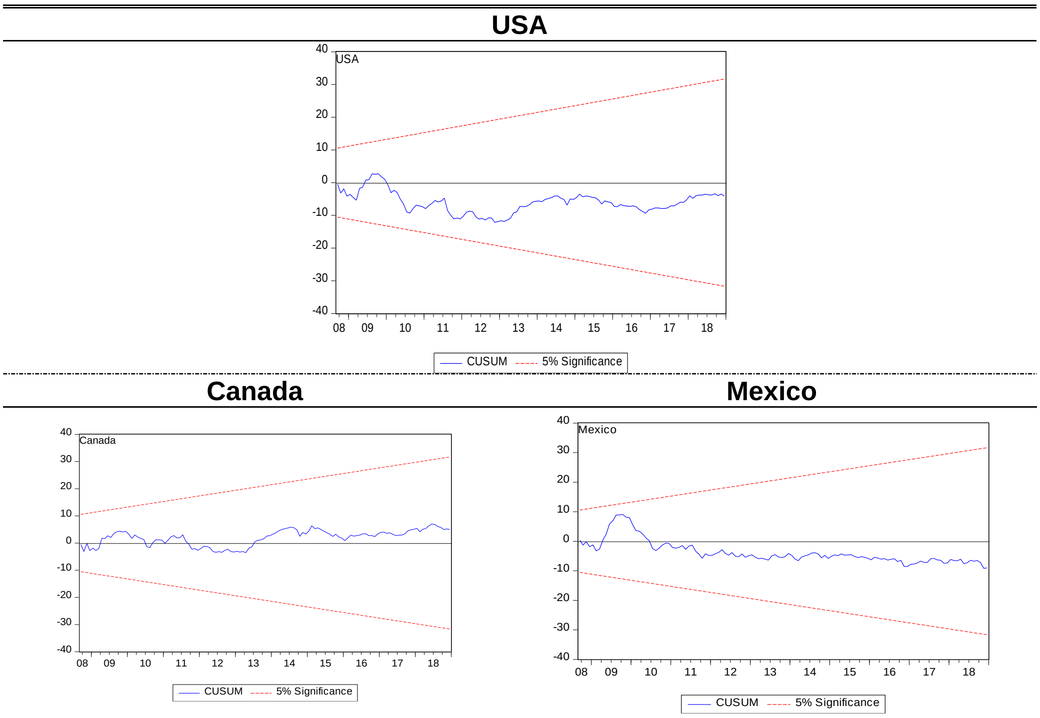


**Figure 1. NAFTA – CuSum test for stability (Sample: 2008-2018)**

**Table 7. NAFTA – Granger causality Test (Sample: 2008-2018)**

| Dependent Variable: D(IndUSA) | | | |
|---|---|---|---|
| Excluded | Chi-Sq | Df. | Prob. |
| D(IndCAN) | 12.205 | 2 | 0.0022 |
| D(IndMEX) | 3.2765 | 2 | 0.1943 |
| **Dependent Variable: D(IndCAN)** | | | |
| Excluded | Chi-Sq | Df. | Prob. |
| D(IndUSA) | 9.8318 | 2 | 0.0073 |
| D(IndMEX) | 4.6667 | 2 | 0.0970 |
| **Dependent Variable: D(IndMEX)** | | | |
| Excluded | Chi-Sq | Df. | Prob. |
| D(IndUSA) | 12.509 | 2 | 0.0019 |
| D(IndCAN) | 10.747 | 2 | 0.0046 |

Granger Causality test is also conducted to further analyze the short-term relations and the direction of causality amongst NAFTA stock markets. According to Granger (1969), a variable Y is said to be Granger-caused by another variable X if X helps in the prediction of Y, or equivalently if the coefficient on the lagged X is statistically significant. The Granger-causality test is implemented in this study to test the presence of leading market(s) amongst the NAFTA markets. The results are reported in Table 7 for the period 2008-2018. The results indicate bidirectional causality (at the 5% level) for the USA and Canadian indices, as they Granger-cause each other. For the USA and Mexican markets, the direction of causality is from the USA to Mexico; there is no reverse causation from Mexico to the USA. This unidirectional pattern holds for the Canadian and Mexican markets as well. Canada Granger-causes Mexico but no reverse causality exists. These results show that the USA and Canadian stock markets can be considered as the leaders within the NAFTA region. The results also once again confirm the reported cointegrating NAFTA market conclusion. If two variables are cointegrated, there must be a causal relationship at least in one direction between them. This is verified with our Granger causality results.

## 5. Conclusion

This paper investigated the integration of the national stock indices of three groups of countries; NAFTA, the Eurozone, and a selected group of OECD countries. We first examine the correlation coefficients within each group. The results show that the correlation coefficients are noticeably higher for the two country groups that are members of a trade or currency zone (NAFTA and the Eurozone) compared to the group of selected OECD countries that are not members of any organization other than being members in OECD. The main objective of this study is to investigate long-term relationships amongst the national stock markets. Given that information provided by correlation coefficients is not enough to claim the existence of market integration, we used the cointegration methodology to study the international integration of markets in the different country groups. The methodology used consisted in applying two alternative tests, the Johansen test and the ARDL bounds test. In the process, we applied the ADF unit root test and found evidence that all the variables in the models are nonstationary, which provides the basis for cointegration analysis. Considering the full samples, the results from the two cointegration techniques unequivocally point to no cointegration in the case of NAFTA, while they are mixed in the case of the other two country groups (Eurozone and the selected OECD countries). The Johansen test found that the markets in those two country groups are cointegrated, while the ADRL bounds test rejects such a conclusion.

To account for possible structural breaks in the data, we broke the analyses down and tested for cointegration in different sub-periods for each country group. The verdicts from the two cointegration techniques were unanimous. The results from both the Johansen and the ARDL bounds testing approaches unveiled evidence of cointegration and long-run relations in each of the sub-periods for each of the three country groups. It can be concluded that financial markets in the OECD countries are cointegrated, and this is true regardless of whether the countries belong to a trade or currency union. These results also confirm the existence of structural breaks and lend support to the dynamic nature of the long-run integration of equity markets and to the argument that markets may exhibit time-varying long-run interdependence. Finally, using the VECM model, the examination of the short-run dynamic relationships amongst the NAFTA stock markets confirmed the cointegration of the markets. The Granger causality test to examine the direction of causality in the NAFTA region indicated bidirectional causality for the USA and Canadian indices. While there is unidirectional causality running from the USA to Mexico on the one hand and from Canada to Mexico on the other, no reverse causality is detected from Mexico to the USA or Canada.

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