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Equity market linkages across Latin American countries

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ABSTRACT

Equity market linkages are of interest to international investors aiming at diversifying their equity portfolio holdings. In fact, the benefit of equity portfolio diversification across different international markets depends on whether markets are integrated or segmented. To discern whether there are any potential benefits to diversification, we investigate the degree of integration across the equity markets of selected Latin American countries (that is, Argentina, Brazil, Chile, Colombia and Peru) by applying dynamic and static cointegration techniques to the largest equity markets of that region. Our aim is to find out whether these equity markets enjoy a long-run relationship as a whole, at the sectoral level, or if they follow different trends. We use weekly equity prices between 2005 and 2023, which marked the end of the Covid-19 pandemic. Our findings suggest that the equity markets are not, as a whole, integrated – with the exception of periods of financial distress. This indicates that there is some potential for international portfolio diversification across Latin American equity markets. On the other hand, our sectoral analysis points to specific diversification opportunities across most of the sectors.

1. Introduction

In contrast to previous decades, in the 1990s Latin American countries (LAC) started a robust process of economic and financial reforms. A significant event was the establishment of Mercosur in 1991, a common market among the largest economies of South America (Argentina, Brazil, Paraguay and Uruguay), joined in later years by other countries (the associate members) of that region. This was followed by financial liberalisation reforms aimed at gradually removing major impediments to cross-border financial transactions, opening domestic banking sectors to foreign banks and allowing foreign investors to operate in domestic equity markets. The liberalisation of domestic equity markets was aimed at providing an additional source of growth financing for domestic companies in addition to the traditional means associated with the domestic banking sector (OECD, 2019). One of the effects of these financial reforms was an increase in correlation among the LAC equity markets whereas, in the past, their markets were more correlated with a dominant external market such as the US (Heaney, Hooper, & Jaugietis, 2002). In 2009, another attempt to further financial integration among LAC was the creation of MILA (Mercado Integrado Latinoamericano), comprising Chile, Colombia, Peru and, later, Mexico (IMF, 2016). Specifically, this was an attempt to integrate the stock market exchanges of these countries (Bolanos, Burneo, Galindo, & Berggrun, 2015).

Given the progress made by Latin American governments to integrate their economies and remove barriers to both trade and finance, in this study we examine the relationship between five leading LAC stock markets to assess whether these reforms have contributed to enhanced financial integration among the equity markets of the LAC. We do this by analysing the long-run relationships

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among their equity markets to discern whether these markets have become more integrated. This is important for both policy makers and investors. The former might see changes in the linkages across equity markets as a means to assess the impact of their reforms, whereas investors across LAC might be interested to know whether equity markets are more integrated and therefore if scope to diversify across LAC is reduced.

We use cointegration techniques to investigate long-run relationships among LAC equity markets. If stock markets are cointegrated, this implies that these markets share a common stochastic trend and therefore any benefit of equity portfolio diversification within the countries would be reduced. Equity markets can still deviate in the short run and temporarily restore the benefits of international portfolio diversification. However, they are expected to be reversed in the long run (Chen, Firth, & Rui, 2002).

Our analysis focuses on the long-run relationship between national equity market indices as well as between sectoral-level equity market indices in the period from July 2005 up to May 2023 (the end of the Covid-19 pandemic). Our study extends prior research focusing on South American equity markets (for example, Bolanos et al., 2015; Chen et al., 2002; Espinosa-Méndez, Gorigoitía, & Vieito, 2017; Mellado & Escobari, 2015; Vides, 2022) in the following two ways. Firstly, the current empirical literature on the financial integration among the Latin American equity markets does not include any studies that investigate the time-varying nature of market integration at sectoral level. In that sense, our study represents a unique contribution to the empirical literature in terms of using sectoral equity indices and dynamic cointegration analysis in the context of LAC over a period including major regional events as well as international shocks like the 2007–2009 global financial crisis and the Covid-19 pandemic. Analysing the relationships among sectoral indices is important as this might address the important question as to whether specific sectors are driving any co-movements towards market integration among LAC stock markets. In addition to this, the analysis of this relationship might help international investors to better spread their equity investments across different sectors of Latin American equity markets: in other words, ascertaining whether industry diversification provides better benefits than country diversification. The second element of novelty of our study, given that the static approach has been widely used in the empirical literature focusing on the Latin -American equity markets (see, for instance, Chen et al., 2002; Yang, Kolari, & Sutanto, 2004; Espinosa-Méndez et al., 2017), is that we investigate the existence (if any) of long-run relationships by adopting a dynamic cointegration approach. As pointed out in several studies (see, for instance, Kearney & Lucey, 2004), a major element of weakness in using static cointegration methodologies is that they miss the element of time variation in equity risk premia. It has been demonstrated that the risk premium of equity is indeed time-varying (see, for instance, Campbell, 1987; Harvey, 1989; Bekaert & Harvey, 1995). Therefore, investigating the integration of equity markets by excluding the possibility of time variation in equity premia may yield unclear and partial results (Kearney & Lucey, 2004). To address this problem, dynamic cointegration methodologies have been used in investigating equity market integration focusing mainly on European markets (see, for instance, Rangvid, 2001; Lucey, Aggarwal, & Muckley, 2004).

The results of the dynamic cointegration approach presented in our study provide evidence of intermittent periods of co-movement at both national and sectoral level over the period 2005–2023. On the other hand, the results of the static cointegration analysis show that there is no evidence of long-run relationships among the LAC equity markets. Overall, our findings show that there is scope for international investors to diversify their equity market portfolios across the five Latin American equity markets.

The rest of our study is organised as follows. Section 2 presents a review of the relevant empirical literature; Section 3 provides a brief overview of major Latin American equity markets; Section 4 describes the methodology and Section 5 presents the data. Empirical results are reported in Section 6 while robustness issues are examined in Section 7. Section 8 summarises and concludes.

2. Related literature

Over the past two decades, the large and rapid expansion of equity markets has drawn significant attention. One of the aspects investigated has been the existence of either co-movements or long-run relationships among equity markets. In this section we classify the recent empirical literature under two groups: the first covering the integration across equity markets located in non-LAC and the second focusing on the degree of integration across the LAC.

2.1. International equity markets integration

Co-movements among markets have generally been investigated using time-varying correlation techniques in a bivariate or multivariate setting based on alternative typologies of dynamic conditional correlation GARCH family models (see, for instance, Silvennoinen & Teräsvirta, 2009; Arouri, Bellalah, & Nguyen, 2010; Aslanidis & Savva, 2011; Kenourgios & Samitas, 2011). One of the key messages from this literature is that correlations among major European equity markets no longer respond to market volatility since the adoption of the Euro as legal tender by twelve European states in 2002 (Silvennoinen & Teräsvirta, 2009). Secondly, correlations among less advanced equity markets tend to increase in times of crisis or uncertainty caused by major economic events (Arouri et al., 2010). Thirdly, by looking at aggregate equity markets and industry indices, correlations among the former are higher in comparison to the latter, providing lower portfolio diversification benefit opportunities than industry indices-based portfolios (Aslanidis & Savva, 2011).

On the other hand, existence of long-run relationships among equity markets have generally been investigated using bivariate and multivariate cointegration techniques (see, for instance, Bachman, Choi, Jeon, & Kopecky, 1996; Huang, Yang, & Hu, 2000; Sheng & Tu, 2000; Aloy, Boutahar, Gente, & Peguin-Feissolle, 2013). The key message from this literature is that the reduction of trade barriers, as well as the implementation of financial liberalisation reforms, are found to be major factors explaining the existence of a long-run relationship among equity markets of the seven major industrial (G-7) countries (Bachman et al., 1996). Conversely, restrictions on capital movements among equity markets characterised by geographical proximity still cause an impediment to the existence of a long-run relationship, even if this might bring benefits in terms of international portfolio diversification, as demonstrated in the case of the Greater China region (Huang et al., 2000).

Furthermore, most of the empirical literature on co-movements and long-run relationships has either focused on advanced equity markets (see, for instance, Bachman et al., 1996; Masih & Masih, 1997; Serletis & King, 1997; Kanas, 1998; Pascual, 2003; Fraser & Oyefeso, 2005; Barari, Lucey, & Voronkova, 2008), or between advanced and emerging markets mainly located in Asia, Africa and eastern Europe (see, for instance, Huang et al., 2000; Sheng & Tu, 2000; Scheicher, 2001; Gilmore & McManus, 2002; Ratanapakorn & Sharma, 2002; Égert & Kočenda, 2007; Li, 2007; Gilmore, Lucey, & McManus, 2008; Olusi & Abdul-Majid, 2008; Syriopoulos & Roumpis, 2009; Savva & Aslanidis, 2010, Guidi, Savva, & Ugur, 2016). The findings related to advanced equity markets highlighted several aspects of the aforementioned long-run relationships. Firstly, factors that explain the presence of a long-run relationship among advanced equity markets are technological change, financial deregulation and international capital-goods trade (see, for instance, Bachman et al., 1996). Secondly, the existence of cointegration (or financial integration) is usually found among advanced equity markets (see, for instance, Masih & Masih, 1997, Serletis & King, 1997) – although it might disappear and then re-emerge by displaying a time-varying path (Fraser & Oyefeso, 2005). On the other hand, the empirical literature focusing on the existence of long-run relationships among advanced and emerging markets presents a different picture. The lack of a common trend among advanced and emerging markets is found in several studies focusing on Pacific-region equity markets and using either the US, Japan or Hong Kong as proxies of advanced markets (see, for instance, Huang et al., 2000; Sheng & Tu, 2000). Other studies focusing on the linkages among European, Middle Eastern and North African equity markets found no evidence of long-run relationships (Olusi & Abdul-Majid, 2008). Furthermore, studies focusing on countries located in Europe show either no long-run relationship (Gilmore & McManus, 2003) or limited evidence of cointegration between pairs of advanced and emerging equity markets (Égert & Kočenda, 2007; Gilmore et al., 2008). Interestingly, intermittent evidence of long-run relationships is found among these markets when a dynamic cointegration approach is used to investigate the time-varying nature (if any) of the long-run relationship between advanced Western Europe and emerging Central Europe equity markets (Gilmore et al., 2008).

2.2. Evidence for Latin America

As reported in Table 1, only a few studies have looked at co-movements or long-run relationships among advanced equity markets and emerging equity markets located in South America, or South American markets only. The findings of these studies very often depend on the period of analysis investigated, the frequency of data as well as the type of methodology used. For instance, Chaudhuri (1997) investigated whether a long-run relationship existed between six South American equity markets (Argentina, Brazil, Chile, Colombia, Mexico and Venezuela) over the period 1985 to 1993. By using the Engle and Granger (1987) cointegration test in a bivariate setting, Chaudhuri (1997) found evidence of long-run relationships between these South American stock markets. Similarly, using bivariate and multivariate cointegration techniques, Chen et al. (2002) analysed the long-run relationship between the same six Latin American stock markets over the period 1995–2000. By dividing the period of analysis into several sub-periods identified in

Table 1
Recent empirical literature on the integration of South American equity markets.

Author(s)	Country(s)	Methodology	Data frequency	Period	Results
Chen et al. (2002)	Argentina, Brazil, Chile, Colombia, Mexico, and Venezuela	Johansen cointegration test	Daily	1995–2000	Equity market indices of the six South American countries share a long-run relationship.
Yang et al. (2004)	Emerging markets (including Argentina, Brazil, Chile, Colombia, Mexico and Venezuela) and the USA	Johansen cointegration test in a bivariate and multivariate setting; Recursive cointegration approach	Monthly	1976–2001	Equity markets of Argentina, Brazil, Chile, and Mexico share a long-run relationship with the USA equity markets during the 1997 Asian financial crisis.
Fernández and Sosvilla-Rivero (2006)	Argentina, Brazil, Chile, Mexico, Peru, and Venezuela and the USA	Johansen cointegration test. Gregory and Hansen cointegration test with structural break.	Daily	1995–2002	Argentina, Chile, and Venezuela equity markets share a long-run relationship with the USA equity market after the 1997 Asian financial crisis.
Diamandis (2009)	Argentina, Brazil, Chile, Mexico and the USA	Johansen and Juselius multivariate cointegration methodology	Weekly	1988–2006	Argentina, Brazil, Chile, Mexico and the USA equity markets are partially integrated.
Dima et al. (2015)	Brazil, Chile and Mexico	Wavelet analysis	Daily	2003–2014	Argentina, Brazil, and Mexico show evidence of cyclical synchronization of their equity markets.
Espinosa-Méndez et al. (2017)	Chile, Colombia, Peru, Mexico and the USA	Dynamic conditional correlation analysis; Gregory-Hansen and Johansen cointegration tests.	Daily	2002–2016	The levels of correlation among the selected Latin American countries increased after the creation of the MILA.
Vatsa et al. (2022)	Argentina, Brazil, Chile, Mexico and USA	Hamilton filter methodology	Weekly	1990–2020	The USA stock market is correlated with the markets of Argentina, Brazil and Mexico, whereas the US and Chilean markets are uncorrelated.

accordance with several global and regional financial crises Chen et al. (2002) found that South American stock markets shared a long-run relationship up until 1999. After 1999, however, no evidence of any long-run relationships was found. From a different perspective, Yang et al. (2004) examined the linkages between 13 emerging stock markets (including Argentine, Brazil, Chile, Colombia, Mexico and Venezuela) and the US stock market during the period from 1976 to 2001. By focusing on the long-run relationship between the US and Latin American equity markets, their findings show that no long-run relationship exists between these markets apart from Argentina. The robustness of the findings is then tested by using a recursive cointegration approach: the results show no evidence of a long-run relationship cointegration until the beginning of the 1997 Asian financial crisis. Therefore, the assumption that long-run relationships (or lack of) among equity markets are stable is clearly challenged by the findings presented in Yang et al. (2004). In a similar study, Hunter (2006) looked at the integration between three Latin American stock markets and the international capital market proxied by the US market over the period 1992–1999. Their findings showed that Argentinean and Chilean stock markets were integrated with the US, whereas no evidence was found in terms of integration between the Mexican and the US markets. More recently, Fernández and Sosvilla-Rivero (2006) examined the linkages between the US and six major Latin American stock markets (Argentina, Brazil, Chile, Mexico, Peru and Venezuela) covering the period 1995–2002. By using the Johansen cointegration test, as well as the Gregory and Hansen (1996) cointegration test with structural break, Fernández and Sosvilla-Rivero (2006) found that that the Johansen cointegration detected a long-run relationship between the US and two Latin American stock markets while the Gregory and Hansen test found long-run relationships with structural break between the US and four Latin American stock markets. Likewise, Diamandis (2009) explored the long-run relationships among the US market and four Latin American stock markets (Argentina, Brazil, Chile and Mexico). By using the Johansen and Juselius (1992) multivariate cointegration methodology, Diamandis (2009) found that the US and these four South American equity markets shared a long-run relationship over the period from 1988 to 2006. Dima, Dima, and Barna (2015) investigated if Brazil, Chile and Mexico showed any evidence of the synchronization of their equity markets over the period 2003–2014. Their findings indicated evidence of cyclical synchronization which was likely due to the effects induced by the creation of the MILA. However, Dima et al. (2015) reported also that the effects of the 2007–2009 global financial crisis disrupted the synchronization among the investigated equity markets and impeded its full restoration over the post-financial crisis period. Espinosa-Méndez et al. (2017) investigated the levels of correlation among LAC equity markets in the period before and after the creation of the MILA. Their findings show that correlation levels in stock returns of member countries increased in the post-MILA period. More recently, Vatsa, Basnet, and Mixon (2022) investigated the degree of interconnectedness among four major Latin American stock markets (that is, Argentina, Brazil, Chile and Mexico) with the US market. Their results show that Latin American markets exhibit strong interlinkages, whereas all of them but Chile are correlated with the US markets. By using the Hamilton filter methodology, Vatsa et al. (2022) also demonstrate that the US can be used as a leading indicator of the cyclical development of the Argentinean stock market.

The empirical literature focusing on the integration among the equity markets of LAC does reveal three important aspects. Firstly, integration among LAC equity markets, as well as with major equity markets like the US, cannot be considered as a permanent condition once detected: in other words, a reverting process might be at work in a period following the detection of integration. Secondly, cointegration tests that might detect a structural break in the long-run relationship among equity markets could explain episodes of intermittent financial integration among the LAC equity markets under study. Thirdly, significant economic events, either at domestic or international level, might either ignite or contribute to the dissolution of long-run relationships among markets. In light of these findings, our study expands upon the geographic perspective of previous empirical studies assessing long-run relationships among South American stock markets by also examining relationships across industries of the same region. Our paper provides insights into international diversification, exploring two key strategies for equity investors: incorporating benchmark market indices for broad exposure or focusing on sectoral market indices to capitalise on industry-specific trends. The choice depends on risk tolerance, return objectives and market outlook.

3. Development of the LAC equity markets

By the end of 2020, the market capitalisation of the Brazilian stock market was the largest in absolute values ($988 bln) out of the South American stock markets, followed by Chile ($184 bln), Colombia ($106 bln) and Peru ($87 bln) (World Bank, 2022).

As shown in Fig. 1, despite a steady decline in the market capitalisation ratio from an average of 120 % (2006–2010) to 84 % (2016–2020), Chile has led in terms of market capitalisation as a percentage of GDP. Relative to the other countries, the market capitalisation ratio for Brazil has shown the sharpest increase in its 5-year average during the last decade, increasing from an average of 40 % (2011–2015) to 54 % in the following years. Colombia has shown a decline in market capitalisation, especially during the period 2016–2020. Like Colombia, the market capitalisation of the Peruvian stock market particularly declined between 2011 and 2015, although it recovered over the following 2016–2020 period. Among South American stock markets, Argentina is far behind, with an average of 11 % of GDP during 2016–2020 confirming it to be the smallest market in terms of capitalisation with respect to GDP.

The number of Latin American domestically listed companies in the five South American stock markets has declined from 996 in 2006 to 908 as of 2020 (see Fig. 2). Brazil is the South American leader with an average number of listed companies far larger than any other individual South American stock markets for each of the five-year average periods. Argentina and Colombia lag well behind the region's leaders, Brazil and Chile, whereas Peru is the only country that experienced increasing numbers of listed companies over the five-year average period.

South American countries introduced reforms aimed at opening their equity markets to foreign investors in the late 1980s or early 1990s (Bekaert, Harvey, & Lundblad, 2003). Table 2 shows three measures of open market measures that have been implemented since then. The first one is the official market liberalisation date itself – that is, the date of formal regulatory change after which foreign

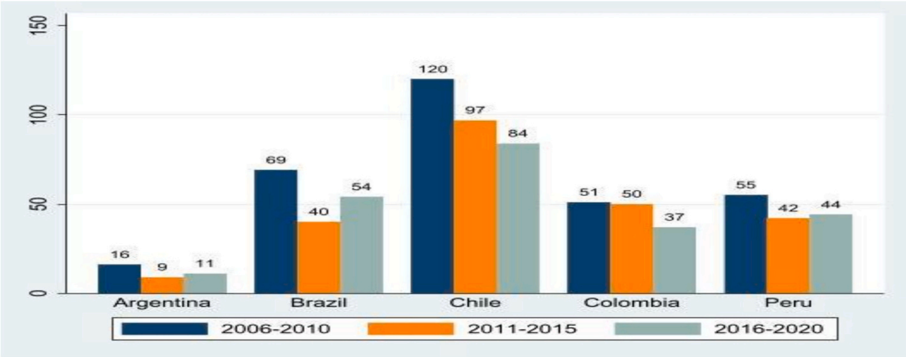

Fig. 1. Market capitalisation of listed companies (% of GDP), 5-year average 2006–2020.
Source: Authors' own calculation on World Bank (2022) data.

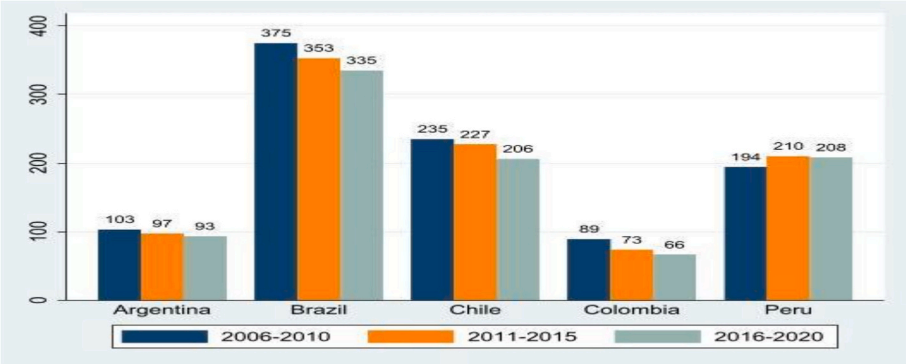

Fig. 2. Listed domestic companies, 5-year average 2006–2020.
Source: Authors' own calculation on World Bank (2022) data.

investors could officially invest in domestic equity securities and domestic investors gained the right to transact in foreign equity securities abroad. Secondly, US banks were allowed to issue American Depository Receipts (ADR) related to publicly listed South American companies. Therefore, international investors could trade these financial instruments, representing shares in foreign stocks, in the US financial markets. This enabled South American public firms to attract foreign investors without the difficulty and expense of listing on the US stock exchange. Another measure is the percentage of foreign country funds invested in the stocks of Latin American companies.

The measures aimed at opening the equity markets and attracting foreign investors had mixed effects. For instance, over the period 2000–2010 the annual portfolio equity net inflows[1] were, on average, the largest in Brazil ($3.61 bln) and Chile ($498 mln), quite modest in in Colombia ($16.8 mln) and in Argentina ($15.5 mln) and negative in Peru ($-69.3 mln).

4. Methodology

To consider the dynamic process of the linkages among South American equity markets, we used the recursive cointegration methods of Hansen and Johansen (1992). Moreover, to overcome the statistical limitation associated with estimating a model with an increasing number of observations, we also used a rolling window variation of the recursive method where the length of the estimation window is fixed. Both these approaches were carried out to test for the presence of dynamic cointegration.

The cointegration analysis, as performed with the Johansen cointegration test, assumes the constancy of the parameters in the VAR models. One of the criticisms of this assumption is that the longer the period of analysis, the more difficult it is to assume that the parameters of the VAR model remain constant. In other words, the results of the Johansen cointegration test do not convey any specific information regarding the temporal stability of the VAR parameters. To overcome this shortcoming, we used a multivariate recursive cointegration methodology proposed by Hansen and Johansen (1992). Their methodology allows us to test the parameter constancy by

[1] As per the definition of World Bank (2022), portfolio equity includes net inflows from equity securities other than those recorded as direct investment and including shares, stocks, depository receipts (American or global) and direct purchases of shares in local stock markets by foreign investors.

Table 2
Equity market opening in major Latin America countries.

Country	Official liberalisation date	First ADR introduction	First country fund introduction
Argentina	November 1989	August 1991	October 1991
Brazil	May 1991	January 1992	October 1987
Chile	January 1992	March 1990	September 1989
Colombia	February 1991	March 1990	September 1989

Source: Bekaert et al. (2003).

using the recursive estimation in three alternative ways: forward recursions, backward recursions and windows of fixed length. Following Hansen and Johansen (1992)'s suggestions, as well as the empirical literature (see, for instance, Yang, Khan, & Pointer, 2003; Phengpis & Apilado, 2004; Barari et al., 2008; Gilmore et al., 2008), we used the forward recursion modality. This recursive approach requires starting the analysis with an initial period from t_0 to t_n with the aim of estimating a VAR model and calculating both λ_{trace} and λ_{max} test statistics. The initial period is then extended by additional j periods and both the λ_{trace} and λ_{max} statistics are re-estimated from t_0 to t_{n+j}. This exercise of extending the initial period carries on until the end of the data is reached. The next step is to then plot graphically both the λ_{trace} and λ_{max} test statistics (calculated as previously explained). In accordance with this procedure, we consider an initial period of three years, which is then extended by an additional period of length equal to one year.[2] To facilitate the interpretation of the results, the calculated λ_{trace} and λ_{max} statistics are rescaled with respect to critical values: if the rescaled values of these statistics are greater than 1.0 then this indicates the presence of cointegration, while values below 1.0 of the rescaled λ_{trace} and λ_{max} indicate no cointegration. However, the findings of the recursive cointegration analysis must be interpreted cautiously. As pointed out by Pascual (2003), in the recursive approach the length of the window increases as new observations are added. This makes the recursive window cointegration test less powerful than a rolling window approach in which the length of the estimation window does not change and rolls forward over time. Therefore, to enrich our analysis, we also carried out a rolling window cointegration approach. To this end, we chose a rolling window size (m), which is the number of consecutive observations for rolling windows equivalent to three years of data ($m = 3$). The number of increments between successive rolling windows is one year, so the first rolling window contains observations from t_o to t_m, the second from t_1 to t_{m+1} and so on. The cointegration analysis was then performed on each rolling window subsample, meaning that for each rolling window we tested for the presence of cointegration by calculating both λ_{trace} and λ_{max} statistics. The calculated statistics were then rescaled similarly to the recursive cointegration analysis, and the results were plotted in order to show graphically how these rescaled values changed over time.

5. Data and variables

In this study we focus on the most important Latin America equity markets in terms of capitalisation: Argentina, Brazil, Chile, Colombia and Peru. The first two are the two largest economies among the founder members of the Mercosur; the last three are the largest economies among the associate Mercosur members.[3] These countries have substantial cross-country investment, close economic ties and similar commodity exports. Their equity markets are also similar in terms of openness and accessibility to foreign investors.[4]

To carry out our econometric analysis, we use both national and sectoral indices for each of our five LAC. The national indices used are: MERVAL (Argentina),[5] IBOVESPA (Brazil),[6] COLCAP (Colombia),[7] IGPA (Chile)[8] and IGBVL (Peru).[9] We use five major sectoral indices for our sectoral analysis of each country, details of which will be explained later. Finally, all equity index prices are expressed in a common currency (US dollar), as the database we used to gather data (Eikon Refinitiv) allows the conversion of the price indices of the selected Latin American equity markets into US dollar with the DataStream exchange conversion facility.

The period of analysis is from 1 July 2005 to 26 May 2023 – a total of 935 weeks. We chose July 2005 as the starting point for our data collection because this is the earliest date at which complete and reliable data for all indices of interest is available. To also include the effect of the Covid-19 pandemic, we chose the end period of May 2023, when the head of the World Health Organization declared an end to Covid-19 as a public health emergency.[10] We use weekly data because weekly information is characterised by less noise, and it minimises the problem that occurs when some markets are closed while others are still open.

Table 3 presents descriptive statistics for the equity market returns calculated as $r_t = [ln(P_t) - ln(P_{t-1})] \times 100$. Average weekly

[2] The reason for choosing an initial period of three years was motivated by the fact that we wanted to start the dynamic analysis with an initial time window that excluded the turmoil experienced by the financial markets following the bankruptcy of Lehman Brothers on September 15, 2008.

[3] These countries became associate Mercosur members in 1996 (Chile), 2003 (Peru) and 2004 (Colombia).

[4] Despite some remaining restrictions, the availability of financial instruments such as American Depositary Receipts and Country Funds enable international investors to easily access those markets (Diamandis, 2009).

[5] The MERVAL is a price-weighted index comprising 28 companies in the Buenos Aires Stock Exchange.

[6] The IBOVESPA is an equity index weighted by trade volumes and it includes the most liquid stocks in the Brazilian equity market.

[7] The COLCAP tracks the performance of the 20 most actively traded shares in the Colombian market.

[8] The IGPA is a capitalisation-weighted index of the major stocks in the Santiago Stock Exchange.

[9] The IGBVL represents the performance of the 29 most actively traded companies in the Lima Stock exchange.

[10] We thank one of the reviewers to point out the importance of including also the Covid-19 period in our empirical analysis.

returns were positive for all stock markets over the entire period except for in Colombia. The Peruvian stock market had the highest weekly average returns (0.165 %) whereas the Colombian stock market had the lowest and negative average weekly returns (−0.129 %). Stock market returns in Argentina were the most volatile in terms of the largest standard deviation (5.569 %), whereas returns from the Chilean stock market were the least volatile with a standard deviation of 3.443 %. Skewness was negative for all equity market returns, indicating that large positive stock returns are more common than negative returns. Table 3 also shows that for each of the LAC stock markets the value of the kurtosis was greater than three, indicating that all returns are leptokurtic having significantly fatter tails and higher peaks compared to the normal distribution. Finally, standard unit root tests such as ADF, PP and KPSS show that the null hypothesis that the returns of the LAC stock market indices are not stationary is rejected at the usual level of significance.

Table 4 shows pairwise correlation coefficients between the weekly returns of the LAC equity markets. The lowest correlation (0.340) is between the Argentinean and Colombian equity markets, whereas the highest (0.628) is between the Chilean and the Brazilian markets.

For the sectoral-level analysis, our dataset comprised weekly equity prices across five different sectors: basic materials, consumer cyclicals, energy, financials and industrials. As stated in McKinsey and Company (2019), these are the most important sectors in the LAC in terms of productivity. Analysing these five sectors is enough to effectively capture the equity portfolio diversification opportunities of the region, while keeping the computational burden within reasonable limits. The composition of each sector follows the structure defined by Thomson Reuters Business Classification Standards.[11] Descriptive statistics of these sectoral indices are presented in Table 5 where several aspects can be highlighted. Firstly, there are five sectors for each country for a total of 25 stock market sectoral indices and the number of observations for each sectoral index is 935. Secondly, descriptive statistics show that the rate of return for sectoral indices of Brazil (Panel B) enjoyed positive average weekly returns over the entire period of analysis for all the five sectoral indices. The rate of returns for Chile (Panel C) and Colombia (Panel D), however, were the ones with the largest number of sector indices (two out of five) with negative average weekly returns over the period. The volatility of the rate of returns, proxied by their standard deviation, was particularly high for the sectoral indices of Argentina and Brazil.

Fig. 3 presents the plots of the five stock market indices as well as their returns over the period 2005–2023. We can observe first that most of the indices showed an upward trend that abruptly came to an end following the beginning of the 2007–2009 global financial crisis. A return to that upward trend is then evident since the end of the global financial crisis. However, the subsequent end of commodities boom that affected major exporters of raw materials like Brazil, Chile and Peru, contributed to the decline of these equity markets before the end of the the first part of the 2010s.

6. Empirical results

6.1. Dynamic cointegration analysis

In this section we present the results of both the recursive and rolling window dynamic cointegration techniques as outlined in Hansen and Johansen (1992). In accordance with those techniques, we calculated the λ_{trace} and λ_{max} statistics and rescaled them to a 5 % critical value before plotting the values for ease of interpretation. As pointed out by Yang et al. (2003), the finding of both recursive and rolling window cointegration techniques can be more informative in comparison to the standard Johansen (1988) and Johansen and Juselius (1990) techniques, as the former technique is able to show any evolving patterns in the long-run relationship.

6.1.1. Recursive cointegration results

Fig. 4 presents the plot of the rescaled recursive λ_{trace} and λ_{max} statistics for the null hypothesis of no cointegration among the LAC equity markets against the alternative hypothesis of at most one cointegration vector. Like in Johansen (1988) and the Johansen and Juselius (1990) multivariate cointegration technique, the recursive cointegration is based on a VAR approach. To this end, one lag was selected in accordance with the Akaike Information Criterion (AIC). We then used two alternative models: one with a constant and trend, another based on a constant only.[12] The time-varying nature of the long-run relationship presented in Fig. 4 shows episodic evidence of cointegration taking place in 2011 as highlighted by the λ_{trace} statistics, whereas there is no evidence of cointegration for the λ_{max} statistics for the whole period of analysis. Therefore, our findings are consistent with the hypothesis of Bekaert and Harvey (2000) that financial liberalisation reforms might not have a significant impact in terms of establishing financial integration among the countries these reforms were implemented in. In other words, equity markets remain segmented even though foreign investors had relatively free access to these markets. In accordance with the recursive cointegration results, the lack of long-run relationships among the LAC equity markets may lead foreign investors to achieve international portfolio diversification benefits by including LAC equity market stocks in their portfolios, thereby obtaining better risk-return trade-offs.

[11] To build these sectoral indices, we utilise the User Customised Index (UCI) tool from Refinitiv Eikon DataStream. As a first step, companies were selected from the Thomson Reuters Eikon Database using market capitalisation, liquidity and sector classification as main criteria. As a result, we obtain a sample of 429 actively traded stocks which have a market capitalisation of at least 70 % of the total, by the end of the year, for each market. After selecting the constituent stocks, sectoral indices are calculated using the market value weighting approach of the DataStream UCI tool. This weighting method eliminates the effects of the exchange rate by applying a common reference currency. From the 429 stocks in the sample, one index was built for each sector and for each country. In total we have twenty-five sectoral indices, five for each of the five LAC of our study.

[12] Results are similar for both models. In our study we present the results of the constant and trend model, whereas the findings of the model with constant only are available upon author request.

Table 3
Descriptive statistics for equity market returns, 2005–2023.

	MERVAL	IBOVESPA	IGPA	COLCAP	IGBVL
Mean	0.120	0.077	0.083	−0.129	0.165
Standard Dev	5.569	5.129	3.443	4.066	3.879
Maximum	19.559	23.386	15.087	24.778	19.134
Minimum	−56.977	−34.515	−28.998	−36.165	−36.292
Skewness	−2.130	−0.568	−1.114	−1.035	−0.923
Kurtosis	20.144	8.043	11.670	15.299	14.410
ADF test	−30.714	−19.751	−31.487	−29.286	−18.442
PP test	−30.772	−32.012	−31.499	−29.322	−27.534
KPSS test	0.041	0.120	0.144	0.095	0.269

Notes: For the Augmented Dickey-Fuller (ADF) and Phillips-Perron tests, the critical values are at −3.43, −2.86, and − 2.57 at the 1 %, 5 %, and 10 % level of significance. If either ADF or PP test statistics are smaller than any of these critical values, the null hypothesis of a unit root is rejected, therefore indicating that the series is stationary. For the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test, the critical values are at 0.739, 0.463, and 0.347 at the 1 %, 5 %, and 10 % level of significance. If the LM statistics of the KPSS test is greater than the critical value, the null hypothesis is rejected (i.e. the series is non-stationary).

Table 4
Correlation coefficients between equity market returns, 2005–2023.

	MERVAL	IBOVESPA	IGPA	COLCAP	IGBVL
MERVAL	1.00				
IBOVESPA	0.561	1.00			
IGPA	0.449	0.628	1.00		
COLCAP	0.340	0.512	0.525	1.00	
IGBVL	0.447	0.559	0.546	0.440	1.00

By grouping similar industries of different LAC countries, we extended the recursive cointegration analysis to the LAC sectoral stock indices and present our findings in Fig. 5. For the basic materials sector equity indices, both the normalised λ_{trace} and λ_{max} statistics (Panels A and B of Fig. 5) do not show any evidence of long-run relationships over the entire period of analysis. Moving to the consumer cyclical sector equity indices, Fig. 5 shows that both normalised λ_{trace} (Panel A) and λ_{max} (Panel B) statistics indicate that there is evidence of temporary episodes of cointegration as both statistics are at times more than the 5 % critical value (above 1.0 in the graph). In particular, the normalised λ_{trace} statistic is above 1.0 in 2009 and 2014, whereas the λ_{max} is above 1.0 in 2011. The findings of the dynamic analysis for the energy sector indices show evidence of periods of intermittent cointegration among the five equities indices: according to the λ_{trace} statistics (Panel A), there is a period of cointegration from 2016 to 2022. These findings are partially confirmed by the λ_{max} that indicates that energy sector indices of the LAC shared a long-run relationship between 2018 and 2021. As for industrial sector indices, the normalised λ_{trace} (Panel A of Fig. 5) indicates evidence of cointegration over two separate periods: the first between 2017 and the beginning of 2018, whereas the second started in the second part of 2019 up to the end of our period of analysis in May 2023. On the other hand, the λ_{max} shows two short-lived periods of cointegration with the first one occurring up to the end of 2008, and the second period between 2017 and 2018. Fig. 5 also shows the time path of λ_{trace} (Panel A) and λ_{max} (Panel B) for the five LAC financial sector equity indices. These financial sector equity indices shared a long-run relationship for several years as the normalised λ_{trace} statistic is above 1.0 from 2011 to 2013, whereas the λ_{max} was above 1.0 from 2011 to 2014.

6.1.2. Rolling windows cointegration results

Fig. 6 presents the rescaled rolling λ_{trace} and λ_{max} statistics for the LAC equity indices when the rolling window approach is applied to test null hypothesis of no cointegration against the alternative of one cointegration vector. The λ_{trace} statistics indicate the presence of cointegration from mid-2010 to the end of 2014 and during 2016. λ_{max} picks up the same initial period ending, however, in late 2013. This approach, unlike the recursive approach, highlights a generally longer period of cointegration among the LAC equity markets following the end of the 2007–2009 global financial crisis. Therefore, the findings presented in Fig. 6 might be interpreted as evidence that diversification benefits at sector level are substantial, except during periods of financial distress as also demonstrated by Attig and Sy (2023).

As in the previous section, we extended the dynamic cointegration analysis to the sectoral stock price indices by grouping similar industries of different LAC countries. Then, using a rolling window approach, we calculated the cointegration statistics and report the findings in Fig. 7. We start with the basic materials sector equity indices, Panel A of Fig. 7 shows that, according to the normalised λ_{trace} statistics, there is some evidence of cointegration from mid-2010 to end of 2012 as well as from 2014 to 2015. For the consumer cyclicals sector equity indices, the results presented in Fig. 7 show that there is evidence of cointegration during the period of the 2007–2009 global financial crisis and its aftermath. Specifically, the normalised λ_{trace} statistics (Panel A) appear to be above 1.0 from the end of 2008 to the end of 2013, and the same outcome can be observed in the case of the λ_{max} (Panel B). Both Panel A and Panel B of Fig. 7 show three periods of an intermittent cointegration relationship among the five energy sector equity indices as indicated by the values above 1.0 of the normalised λ_{trace} (around 2011 and between the end of 2018 and 2019 as shown in Panel A) and λ_{max} statistics

Table 5
Summary statistics of sector indices returns (%).

Panel A: Argentina	Basic materials	Consumer cyclical	Energy	Financials	Industrials
Mean	0.063	−0.015	0.02	0.084	0.097
St. Dev	5.426	4.836	5.677	5.689	5.386
Maximum	21.525	18.016	75.788	24.940	23.081
Minimum	−46.229	−37.402	−36.690	−63.733	−48.946
Skewness	−1.126	−1.341	2.029	−2.192	−1.195
Kurtosis	13.218	14.030	45.586	23.837	14.751
ADF	−30.391	−29.841	−30.578	−30.048	−19.047
PP test	−30.473	−30.048	−31.251	−30.049	−29.886
KPSS test	0.085	0.079	0.104	0.099	0.089
Panel B: Brazil					
Mean	0.05	0.069	0.01	0.088	0.085
St. Dev	5.983	5.468	6.163	5.797	4.937
Maximum	35.908	22.224	26.693	31.398	30.161
Minimum	−30.629	−44.301	−36.526	−33.025	−35.245
Skewness	−0.002	−1.043	−0.560	−0.303	−0.704
Kurtosis	7.299	10.128	7.310	8.048	10.351
ADF	−19.628	−19.088	−29.649	−20.297	−19.998
PP test	−32.073	−30.804	−29.736	−33.016	−33.617
KPSS test	0.122	0.214	0.094	0.118	0.107
Panel C: Chile					
Mean	0.026	−0.022	0.014	0.055	−0.032
St. Dev	3.201	4.485	4.073	3.401	3.652
Maximum	13.154	24.304	15.982	14.119	15.275
Minimum	−24.483	−30.950	−28.496	−33.088	−28.698
Skewness	−0.649	−0.617	−0.699	−1.205	−1.076
Kurtosis	7.95	8.891	8.247	13.553	10.854
ADF	−30.704	−31.459	−30.193	−31.737	−28.692
PP test	−30.705	−31.447	−30.198	−31.738	−28.721
KPSS test	0.126	0.443	0.106	0.222	0.107
Panel D: Colombia					
Mean	−0.031	−0.04	0.058	0.027	0.028
St. dev	4.701	8.927	5.395	3.620	4.204
Maximum	42.364	46.132	27.180	16.092	19.650
Minimum	−38.848	−150.092	−49.514	−25.093	−23.156
Skewness	−0.713	−5.954	−1.098	−1.344	−0.335
Kurtosis	19.541	98.752	14.339	12.337	6.717
ADF	−31.977	−30.465	−28.738	−18.665	−18.833
PP test	−31.944	−30.744	−28.815	−29.390	−30.395
KPSS test	0.426	0.139	0.419	0.489	0.582
Panel E: Peru					
Mean	0.111	0.095	−0.155	0.191	0.128
St. Dev	4.004	3.559	6.029	2.754	4.099
Maximum	13.368	36.426	42.937	15.910	28.992
Minimum	−33.279	−29.797	−37.359	−16.580	−19.498
Skewness	−0.902	1.331	0.386	−0.048	0.360
Kurtosis	10.893	30.871	11.536	7.990	8.348
ADF	−28.622	−25.015	−30.510	−24.822	−26.406
PP test	−29.366	−25.149	−31.052	−25.391	−26.859
KPSS test	0.205	0.503	0.073	0.448	0.606

(in early 2013 and around 2015 and in 2019 as shown in Panel B). These results are partially consistent with the dynamic cointegration recursive approach, though the length of the cointegration periods was much shorter in the case of the dynamic cointegration rolling windows. The financial sector of the LAC (Fig. 7) is also characterised by intermittent periods of cointegration. These periods occurred mainly in 2009, the period from the beginning of 2011 to the second part of the 2013 and between 2017 and 2018 with an additional short-lived period in the second part of 2022 as shown by the normalised λ_{trace} in Panel A of Fig. 7. For λ_{max} (Panel B of Fig. 7), we observe shorter periods of cointegration among the LAC financial sector indices which occurred from mid-2011 to the beginning of 2013. Interestingly, the periods in which the evidence of cointegration was found tend to coincide with major declines of the financial sector equity indices. From 2011 to 2014 most of these financial sector equity indices reported negative returns. The results of the rolling window approach do in some ways differ if compared to the recursive approach in relation to the length of the cointegration

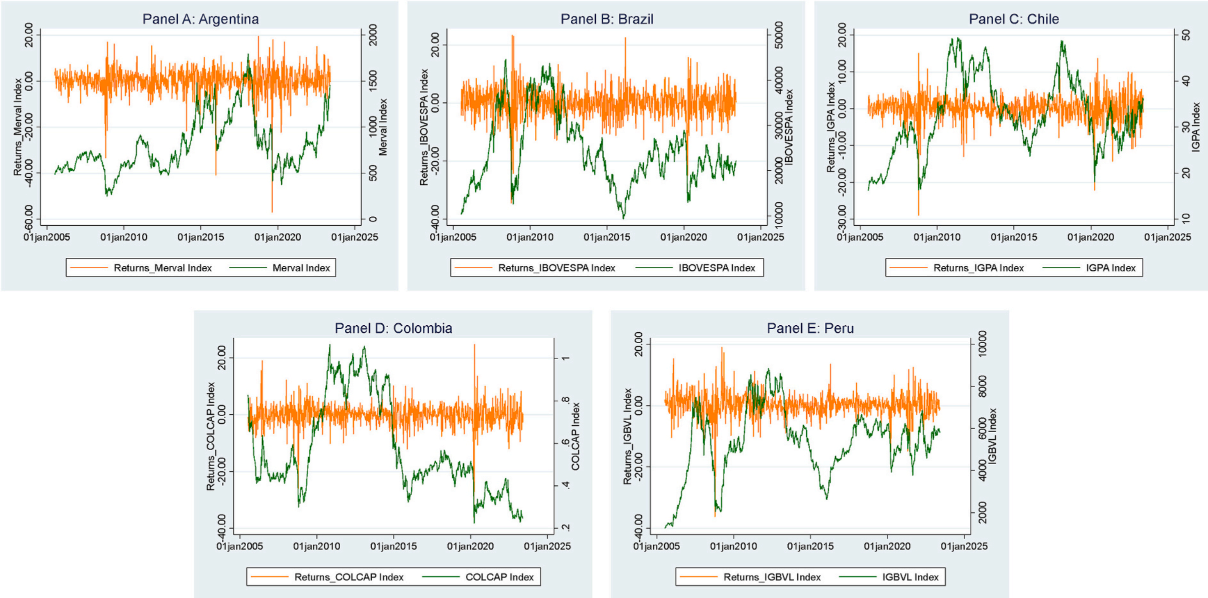

Fig. 3. LAC stock index prices and returns.

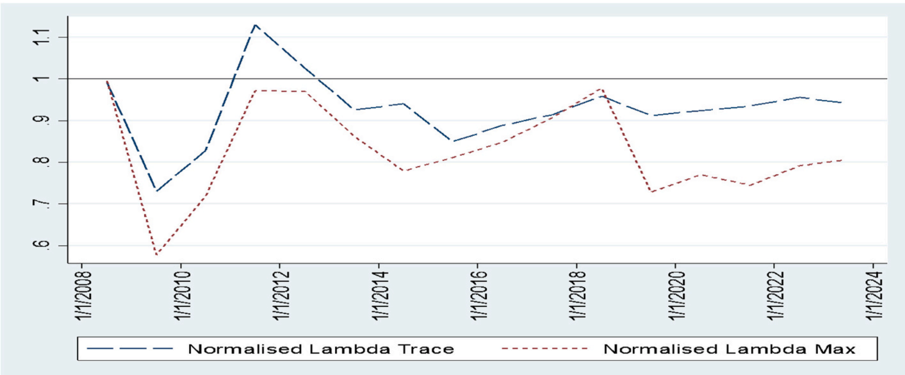

Fig. 4. Recursive normalised test statistics for LAC equity markets.

Notes: This figure illustrates values for the Hansen and Johansen (1992) recursive lambda trace and lambda max statistics for $H_0 : r = 0$ (no cointegration) against $H_1 : r = 1$ (one cointegration relation in the system), rescaled to a 5 % critical value. The value of the statistics (either the normalised lambda trace or the normalised lambda max) above 1.0 indicates the presence of a cointegration relationship, whereas below 1.0 indicates no cointegration.

periods: despite this, however, both approaches tend to confirm evidence of episodic cointegration between the financial sector equity indices of the five LAC. Finally, the normalised λ_{trace} and λ_{max} statistics for the industrials sector show that both λ_{trace} (Panel A of Fig. 7) and λ_{max} (Panel B of Fig. 7) indicate some evidence of cointegration between 2011 and 2012.

7. Robustness checks

In this section we conduct additional analysis to compare the consistency of the dynamic cointegration results when a static approach to the analysis of long-run movements is investigated. Secondly, we tested for the possible presence of cointegration with structural breaks by using alternative methodologies.

7.1. Static cointegration analysis and structural break tests

Using a static multivariate cointegration analysis suggested by Johansen (1988) and Johansen and Juselius (1990), we first tested for the presence of long-run relationships among five leading Latin American equity markets in a multivariate setting. In addition to this, we also used the Gregory and Hansen (1996) cointegration test to detect structural breaks that can reveal evidence of long-run

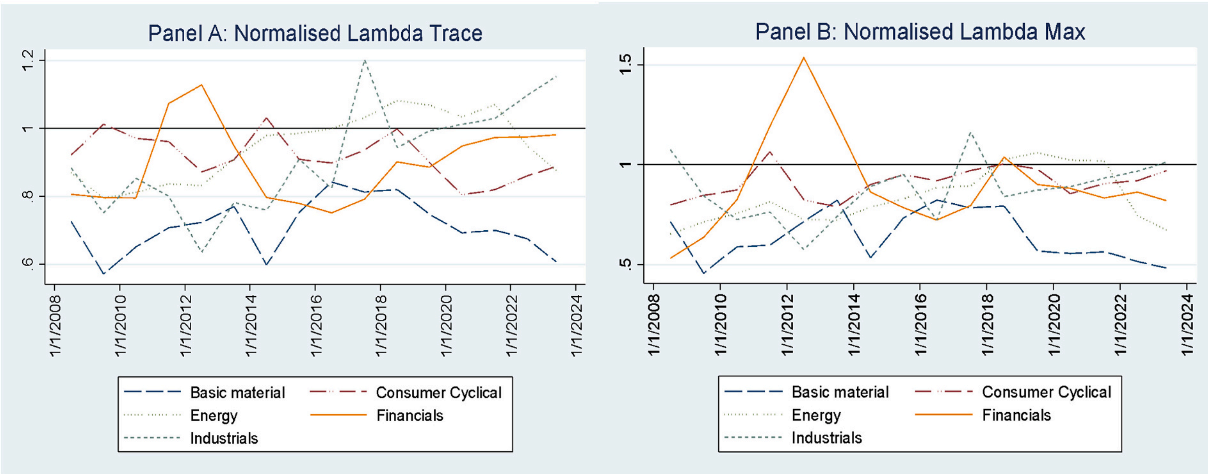

Fig. 5. Recursive normalised test statistics for sectors.
Notes: This Figure illustrates values for the Hansen and Johansen (1992) recursive lambda trace (Panel A) and lambda max (Panel B) statistics for $H_0 : r = 0$ (no cointegration) against $H_1 : r = 1$ (one cointegration relation in the system), rescaled to a 5 % critical value. The value of the statistics (either the normalised lambda trace or the normalised lambda max) above 1.0 – that is, above the horizontal line – indicates the presence of cointegration.

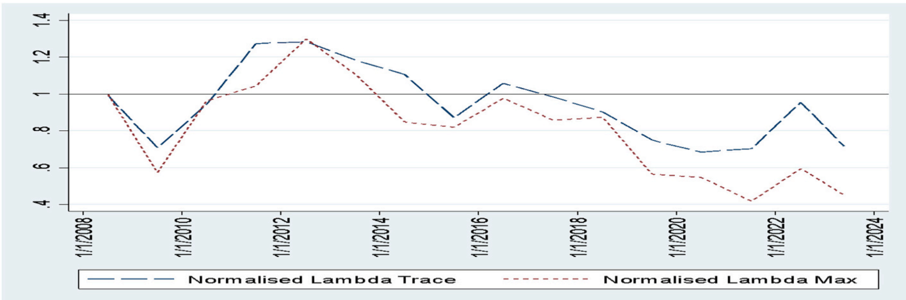

Fig. 6. Rolling windows normalised test statistics for the LAC national equity indices.
Notes: This figure illustrates values of the rolling window lambda trace and lambda max statistics of the Hansen and Johansen (1992) cointegration test for $H_0 : r = 0$ (no cointegration) against $H_1 : r = 1$ (one cointegration relation in the system), rescaled by 5 % critical value. The value of the statistics (either the normalised lambda trace or the normalised lambda max) above 1.0 – that is, above the horizontal line – indicates presence of cointegration.

relationships that are not identified by the standard cointegration methodologies such as the Johansen (1988) and Johansen and Juselius (1990) techniques. Gregory and Hansen (1996) argue that Johansen (1988)'s cointegration test does not consider the possibility that a cointegration vector can change its parameters at an unknown time. In other words, the specific case of testing for cointegration assumes that the cointegrating vector is time-invariant, whilst the more general case assumes that the parameters of the cointegration vectors can change at an unknown date. Gregory and Hansen (1996) suggest three alternative models accommodating changes in the parameters of the cointegrating vector under the alternative of no changes. The first one is the so-called level shift model (model C) that allows for changes in the intercept only. The second model accommodating a trend in data also restricts a shift only to change in level with a trend (model C/T). The third model allows for changes both in the intercept and slope of the cointegrating vector (model C/S). Each of these models therefore allows a structural change in one or more of their parameters. In each of these models, we tested the null hypothesis of no cointegration versus the alternative of cointegration when we take into account a possible regime shift in the LAC equity markets. A rejection of the null hypothesis implies there is a long-run relationship among these markets. To summarise, along with the standard Johansen (1988) cointegration approach, we also implemented Gregory and Hansen (1996)'s cointegration test in this study. We used both these cointegration tests on stock market indices at national level as well as on stock market indices at sectoral level. We applied these tests in a static fashion: in other words, the outcome of the tests would cover the entire period of analysis.

7.1.1. Static multivariate cointegration test results

The Johansen (1988) and Johansen and Juselius (1990) techniques are based on a VAR approach which addresses two main issues.

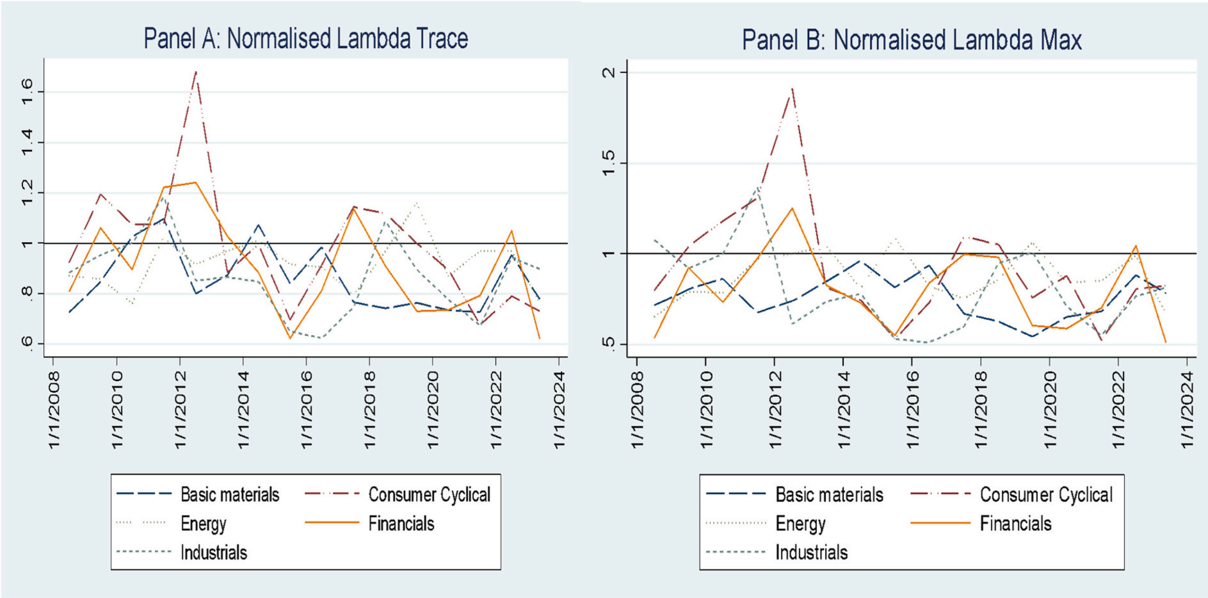

Fig. 7. Rolling windows normalised test statistics for LAC sectoral equity indices.
Notes: This figure illustrates values of the recursive lambda trace (Panel A) and lambda max (Panel B) statistics of the Hansen and Johansen (1992) cointegration test for $H_0: r = 0$ (no cointegration) against $H_1: r = 1$ (one cointegration relation in the system), rescaled to a 5 % critical value. The value of the statistics (either the normalised lambda trace or the normalised lambda max) above one – that is, above the horizontal red line – indicates presence of cointegration.

The first selects an appropriate number of lags in the VAR to eliminate the presence of serial correlation in the residual of the estimated VAR model. We addressed that issue by using the AIC which indicated two lags to remove serial correlation in the residuals of the VAR model for our five LAC equity markets. The second main issue is whether to include a constant and a trend in the VAR model. To this end we estimated two alternative VAR models, one based on the trend and constant version of the Johansen's cointegration model and an additional one with a constant only. Results for both models are similar and we reported only the results of the model with trend and constant. After addressing these two main issues, we then tested for a long-run relationship among our five LAC equity markets via the λ_{trace} and λ_{max} statistics. The findings, presented in Table 6, show that the null hypothesis that the equity markets of the five LAC are not cointegrated ($r = 0$) against the alternative of one cointegrating vector ($r \leq 1$) cannot be rejected because the test statistics, for both λ_{trace} and λ_{max}, do not exceed critical values at the 5 % level of significance. This implies that there is no long-run relationship among the LAC equity markets; that is, they move far apart in the long run. The results of no linkages among the equity markets of LAC can also be interpreted from an international portfolio diversification perspective. In other words, the absence of a long-run relationship implies that international investors might benefit in terms of risk reduction by diversifying their international equity portfolio by investing across the five LAC equity markets.

To the best of our knowledge there are no similar previous studies focusing on the five LAC we are interested in, but we attempted a comparison with related studies focusing on the South American equity markets to check the robustness of our findings. For instance, Chaudhuri (1997) uses monthly data to investigate bivariate cointegration among six South American equity markets. His findings reveal the presence of cointegration between pairs of South American equity markets over the period 1986 to 1993, but no multivariate cointegration test was performed. Diamandis (2009) uses weekly data over the 1988 to 2006 period and applies a multivariate cointegration analysis to five major South American equity markets as well as the US equity market. His findings show evidence of cointegration among these markets. More recently, Vides (2022), in his recent investigation of LAC equity markets via a fractional cointegration approach, concludes that these markets tend to move together in the long run. Therefore, relative to Diamandis (2009) and Vides (2022), our findings at the country level point more strongly towards a lack of long-run relationship among the LAC equity markets.

To explain the cointegration among the equity markets of G-7 countries, Bachman et al. (1996) reported three potential determinants of a long-run relationship among stock prices: international capital-goods trade, financial deregulation and technological change. The first hypothesis would imply that the reallocation of capital-goods, through international trade from one country to another, would begin to equate the marginal product of capital across countries, therefore increasing the efficiency of firms. In our study, two out of five LAC are member states of the Mercosur whereas the three other countries are member associates. Since the Mercosur is a common market that incorporates all aspects of the customs union and extends it by allowing free movements of factors of production, the fact that three countries of our sample are associate members might then have prevented them from enjoying the benefits of a full membership including the trade of capital-goods. The second hypothesis is based on financial deregulation. LAC have taken significant steps to open their capital markets and even created a unified capital market through the MILA. However, it might be

Table 6
Multivariate Johansen cointegration test for LAC stock market indices.

Null hypothesis	Alternative hypothesis	λ_{trace}	CV 5 % (trace)	λ_{max}	CV 5 % (max)
$r = 0$	$r \leq 1$	65.819	69.818	27.268	33.876
$r = 1$	$r \leq 2$	38.550	47.856	19.688	27.584
$r = 2$	$r \leq 3$	18.861	29.797	13.137	21.131
$r = 3$	$r \leq 4$	5.724	15.494	3.898	14.264
$r = 4$	$r \leq 5$	1.825	3.841	1.825	3.841

Notes: The number of cointegrated vectors is indicated by r. CV stands for critical value. ***, **, * indicate rejection of the null hypothesis at the 1 %, 5 % and 10 % level of significance respectively. Results presented in this table are based on the trend and constant version of Johansen's cointegration model. To check the robustness of these results we also performed Johansen's cointegration with the constant only: the findings are similar to the ones presented in this table.

the case that these structural reforms have not yet contributed to create a common trend among the equity prices of the LAC countries, therefore explaining the lack of cointegration among the selected LAC. A final hypothesis is the presence of a dominant economy which could play a role of source of technological change, in other words cointegration among stock markets is driven by a common technological shock (Bachman et al., 1996). Despite being the largest economy among the selected LAC, Brazil still ranks behind in terms of R&D expenditure and export of high-tech products (Esteves & Feldmann, 2016). This might have impeded the Brazilian economy from becoming the principal source of technological change in the Latin America region and therefore playing a leading role as a source of technological change in the region.

For the second part of our static cointegration analysis, we applied the same methodology to each of the five sectors (basic materials, consumer cyclicals, energy, financials and industrials) across the LAC. In other words, we grouped together the same sector equity indices of the five different countries, and, for each group, we tested for the presence (if any) of a long-run relationship. This enabled us to assess whether the same sectors in different countries share the same trend over the entire period of analysis.[13] Our findings are presented in Table 7 which shows the results of the multivariate Johansen cointegration tests at sectoral level. As can be observed, for almost every group of sectors the null hypothesis of no cointegrating vector ($r = 0$) cannot be rejected, implying that there is no long-run relationship within each group of sectors as presented in Table 7. In particular, the values of λ_{trace} and λ_{max} test statistics are not greater than their corresponding critical values at 5 % level of significance for sectors such as basic materials (Panel A), consumer cyclicals (Panel B), energy (Panel C) and financials (Panel D): therefore, these groups of sectoral equity indices do not exhibit any tendency to move together in the long run. Conversely, as Panel E of Table 7 indicates, the industrial sector indices of the five LAC stock markets tend to move in the same direction in the long run as the values of the test statistics for both λ_{trace} and λ_{max} are greater than the corresponding critical values.

The general lack of cointegration among most of the LAC equity indices grouped by sector implies that there are potential gains in risk reduction for international investors willing to build an international equity portfolio made up of the stocks of LAC companies operating in the same sector but in different countries. For example, the lack of a long-run relationship among LAC consumer cyclicals indices implies that an international equity portfolio made up of these equity stocks would aim to provide diversified exposure to LAC cyclical consumer companies.[14]

7.1.2. Cointegration with structural break results

As pointed out previously, Gregory and Hansen (1996)'s cointegration test enables us to detect any structural breaks in the cointegration relationship that are not identified by a conventional cointegration test such as the Johansen test. In this section we present the Gregory and Hansen (1996) cointegration test results in Table 8. The findings are based on a battery of three statistical tests (ADF^*, Z_t^* and Z_α^T) calculated on three alternative structural change models as described in the Appendix A. The null hypothesis for each of the tests is that cointegration does not exist, whereas the alternative hypothesis is that cointegration exists with a structural break at an unknown date. We then take the statistics with the smallest values across all possible break points so that we can infer whether a structural break took place over the period 2005–2023. Panel A of Table 8 reports the Gregory and Hansen (1996) cointegration test results for the LAC stock market indices at the aggregate country level. The values of the statistics (i.e. ADF^*, Z_t^* and Z_α^t) are greater than the 5 % critical values for each of the specification models used. These findings demonstrate that we cannot reject the null hypothesis of no cointegration at the 5 % critical value. This means that LAC equity markets did not share any long-run relationship over the entire period of our analysis. Therefore, the results of the Gregory and Hansen (1996) cointegration test are consistent with the findings of the static cointegration technique presented in the previous sub-section.

We repeated the Gregory and Hansen (1996) cointegration analysis in the five groups of equity market indices at sectoral level. The

[13] The optimal number of lags, as determined by the AIC, was two for all the VAR models based on sectoral indices. Each VAR model for each group of indices was estimated in two alternative ways: with a constant and trend in the cointegrating vector as well as with a constant only in the cointegrating vector.

[14] This is very often the case with sector Exchange Trade Funds (ETF) that invest specifically in the stocks of a particular industry or sector.

Table 7
Multivariate Johansen cointegration results - sector equity indices.

Null hypothesis	Alternative hypothesis	λ_{trace}	CV 5 % (trace)	λ_{max}	CV 5 % (max)
Panel A: Basic Materials					
$r = 0$	$r \leq 1$	42.381	69.818	16.407	33.876
$r = 1$	$r \leq 2$	25.974	47.856	12.679	27.584
$r = 2$	$r \leq 3$	13.294	29.797	7.459	21.131
$r = 3$	$r \leq 4$	5.835	15.494	4.310	14.264
$r = 4$	$r \leq 5$	1.525	3.841	1.525	3.841
Panel B: Consumer cyclicals					
$r = 0$	$r \leq 1$	62.014	69.818	32.886	33.876
$r = 1$	$r \leq 2$	29.127	47.856	15.114	27.584
$r = 2$	$r \leq 3$	14.012	29.797	8.761	21.131
$r = 3$	$r \leq 4$	5.251	15.494	3.566	14.264
$r = 4$	$r \leq 5$	1.685	3.841	1.685	3.841
Panel C: Energy					
$r = 0$	$r \leq 1$	61.221	69.818	22.848	33.876
$r = 1$	$r \leq 2$	38.372	47.856	16.292	27.584
$r = 2$	$r \leq 3$	22.080	29.797	13.399	21.131
$r = 3$	$r \leq 4$	8.680	15.494	5.039	14.264
$r = 4$	$r \leq 5$	3.641	3.841	3.641	3.841
Panel D: Financials					
$r = 0$	$r \leq 1$	68.506	69.818	29.133	33.876
$r = 1$	$r \leq 2$	39.373	47.856	17.229	27.584
$r = 2$	$r \leq 3$	22.143	29.797	12.816	21.131
$r = 3$	$r \leq 4$	9.327	15.494	5.914	14.264
$r = 4$	$r \leq 5$	3.412	3.841	3.412	3.841
Panel E: Industrials					
$r = 0$	$r \leq 1$	80.539*	69.818	34.319*	33.876
$r = 1$	$r \leq 2$	46.220	47.856	25.011	27.584
$r = 2$	$r \leq 3$	21.208	29.797	14.567	21.131
$r = 3$	$r \leq 4$	6.640	15.494	4.777	14.264
$r = 4$	$r \leq 5$	1.863	3.841	1.863	3.841

Notes: The number of cointegrated vectors is indicated by r. CV stands for critical value. ***,**,* indicate rejection of the null hypothesis at the 1 %, 5 % and 10 % level of significance respectively. Results presented in this table are based on the trend and constant version of Johansen's cointegration model. To check the robustness of these results we also performed Johansen's cointegration with the constant only: the findings are similar to the ones presented in this table.

goal was to investigate whether grouped together indices belonging to the same sector shared a long-run relationship with a structural break over the whole period of analysis. Panels B to F of Table 8 show the results for the basic materials, consumer cyclicals, energy, financials and industrials equity indices respectively. Our findings, for all sectors and for all models, show that we cannot reject the null hypothesis of no cointegration at the 5 % critical value for all sectors but industrials. This indicates that, at the sectoral level in general, there is no evidence of a long-run relationship over the entire period of the analysis. There could therefore be benefits to diversifying LAC equity investment portfolios by including stocks of LAC companies operating in the same sectors and located in different countries in South America. Thus, the analysis conducted using the Gregory and Hansen (1996) technique is consistent with the results reported in the previous section using the Johansen (1988) and Johansen and Juselius (1990) techniques.

One of the issues with the Gregory and Hansen (1996) cointegration test is that the test is based on the hypothesis of cointegration with a single structural break and therefore results of the test may be questionable if there is more than one structural break.[15] As pointed out by Bai and Perron (1998), if the true number of breaks is two, then Gregory and Hansen (1996)'s test is misspecified and would perform poorly. To overcome this issue, Bai and Perron (1998) introduced a test allowing for cointegration with an unknown number of breaks. We applied Bai and Perron's test to both national and sectoral indices and found evidence of at least five structural breaks at national and sectoral level.[16] One of the shortcomings of Bai and Perron (1998) is that the test does not include the possibility of considering alternative models that the Gregory and Hansen (1996) test allows for, therefore the Bai and Perron's results might not be comparable with the Gregory and Hansen results we presented earlier. To overcome this disadvantage, we used Maki (2012)'s

[15] We thank one of the anonymous referees for pointing this out.
[16] Results are available upon author request.

Table 8
Multivariate Gregory and Hansen (1996) cointegration test results for LAC stock market indices.

Panel A: National stock market indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−4.69	15/12/2017	−4.40	30/03/2018	−39.77	30/03/2018
Model C/T	−4.95	10/08/2018	−4.61	25/05/2018	−41.89	25/05/2018
Model C/S	−4.82	29/06/2018	−4.72	10/09/2010	−44.13	10/09/2010
Panel B: Basic material indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−3.18	22/06/2018	−2.65	1/06/2018	−17.85	1/06/2018
Model C/T	−3.76	22/06/2018	−3.23	1/06/2018	−25.11	1/06/2018
Model C/S	−4.63	6/07/2018	−3.96	20/7/2018	−33.42	20/07/2018
Panel C: Consumers cyclical indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−4.40	17/10/2014	−3.87	26 Sept 2014	−30.16	26 Sept 2014
Model C/T	−4.76	17/10/2014	−4.32	26 Sept 2014	−36.70	26 Sept 2014
Model C/S	−4.92	17/04/2015	−4.67	3 Apr 2015	−40.83	3 Apr 2015
Panel D: Energy indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−4.41	18/09/2020	−4.40	18/09/2020	−38.83	18/09/2020
Model C/T	−5.15	25/12/2015	−5.11	27/11/2015	−50.71	27/11/2015
Model C/S	−5.76	1/11/2019	−6.25	2/11/2018	−74.16	2/11/2018
Panel E: Financial indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−4.79	9/03/2018	−4.76	6/07/2018	−39.49	6/07/2018
Model C/T	−5.45	29/06/2018	−5.70	6/07/2018	−55.73	6/07/2018
Model C/S	−5.53	28/11/2014	−5.67	6/07/2018	−56.51	6/07/2018
Panel F: Industrials indices						
	ADF^{*}	Break point	Z_{t}^{*}	Break point	Z_{α}^{*}	Break point
Model C	−3.16	1/03/2013	−2.97	1/03/2013	−18.00	1/03/2013
Model C/T	−4.68	26/10/2018	−4.58	26/10/2018	−41.59	26/10/2018
Model C/S	−3.85	10/02/2017	−3.50	3/07/2015	−90.35**	3/07/2015

Notes. ADF^{*}, Z_{t}^{*} *and* Z_{α}^{*} test the null hypothesis of no cointegration against the alternative hypothesis of cointegration with a possible regime shift (also called structural break). For the ADF^{*} test, the 5 % critical value for Model C is −5.56, for Model C/T it is −5.83, whilst for model C/S it is −6.41. For the Z_{t}^{*} test statistic, the 5 % critical value for Model C is −5.56, for Model C/T it is −5.83, whilst for model C/S it is −6.41. For the Z_{α}^{*} test statistics, the 5 % critical value for Model C is −59.4, for Model C/T it is −65.44, whilst for model C/S it is −78.52. The null hypothesis of no cointegration is rejected if the test statistic is smaller than the corresponding 5 % critical value. ** denotes rejection of the null hypothesis of no cointegration at 5 % critical value.

methodology that allows us to test for cointegration with an unknown number of breaks through the following models: level shift model (model C); level shift with trend (model C/T); regime shift model (model C/S); trend and regime shift model. For all these alternative models, the null hypothesis of no cointegration is rejected in favour of the alternative hypothesis of cointegration with an unknown number of breaks if the Maki test statistics is less than a selected critical value. Our results presented in Table 9 show evidence of cointegration with five structural breaks for energy sector equity indices (models C and C/S in Panel D) as well as financial indices (model C/S in Panel E), whereas there is no evidence of cointegration with structural breaks for all other sectors (Panels B, C and F) as well as national equity indices (Panel A). Therefore, the Maki test results are widely consistent with most of the results we reported in Table 8 using the Gregory-Hansen test. The causes of the existence of a long-run relationship with structural breaks among the industrial sector equity indices as well as financial indices as reported in Panels D and E of Table 8, may be quite difficult to identify.

Table 9
Maki (2012) cointegration test results.

Panel A – National stock market indices								
	Maki test statistic	Critical value (5 %)	No of breaks	Break 1	Break 2	Break 3	Break 4	Break 5
Model C	−4.042	−6.306	5	07/09/2012	27/06/2014	04/11/2016	09/08/2019	09/07/2021
Model C/T	−4.160	−6.494	5	22/06/2007	19/09/2014	04/11/2016	09/08/2019	06/08/2021
Model C/S	−7.338	−8.869	5	04/07/2008	02/11/2012	22/08/2014	08/06/2018	06/08/2021
Trend and Regime shifts	−7.388	−9.482	5	29/10/2010	28/01/2013	11/12/2015	09/08/2019	04/06/2021
Panel B: Basic material indices								
Model C	−3.478	−5.992	3	19/09/2014	09/08/2019	06/08/2021	–	–
Model C/T	−5.184	−6.494	5	25/01/2008	28/09/2012	19/09/2014	19/04/2019	26/03/2021
Model C/S	−4.520	−7.244	2	11/12/2015	06/08/2021	–	–	–
Trend and Regime shifts	−9.160	−9.482	5	06/03/2009	16/12/2011	11/12/2015	26/01/2018	13/12/2019
Panel C: Consumers cyclical indices								
Model C	−6.133	−6.306	5	09/11/2007	03/06/2011	29/03/2013	23/10/2015	03/08/2018
Model C/T	−5.238	−6.494	5	25/01/2008	20/11/2009	30/11/2012	23/10/2015	02/02/2018
Model C/S	−8.438	−8.869	5	01/06/2007	22/05/2009	05/09/2014	04/05/2018	12/02/2021
Trend and Regime shifts	−8.655	−9.482	5	20/04/2007	08/06/2012	16/01/2015	14/07/2017	07/06/2019
Panel D: Energy indices								
Model C	−6.316**	−6.306	5	20/02/2009	08/06/2012	11/12/2015	02/08/2019	28/05/2021
Model C/T	−6.295	−6.494	5	04/05/2007	20/02/2009	11/12/2015	22/12/2017	29/11/2019
Model C/S	−9.004**	−8.869	5	20/02/2009	04/11/2011	17/01/2014	11/12/2015	30/07/2021
Trend and Regime shifts	−9.405	−9.482	5	24/08/2007	05/11/2010	17/01/2014	11/12/2015	11/05/2018
Panel E: LAC Financial indices								
Model C	−4.870	−6.306	5	29/10/2010	04/10/2013	07/08/2015	14/07/2017	09/08/2019
Model C/T	−5.191	−6.494	5	10/07/2009	15/06/2012	23/10/2015	13/10/2017	09/08/2019
Model C/S	−9.342**	−8.869	5	14/12/2007	04/05/2012	22/08/2014	11/08/2017	09/08/2019
Trend and Regime shifts	−9.167	−9.482	5	14/12/2007	04/11/2011	29/08/2014	11/08/2017	09/08/2019
Panel F: LAC Industrials								
Model C	−4.023	−6.306	5	01/06/2007	26/06/2009	23/06/2017	09/08/2019	16/07/2021
Model C/T	−5.648	−6.055	2	30/12/2016	06/08/2021	–	–	–
Model C/S	−6.048	−8.869	5	04/05/2007	27/01/2012	11/08/2017	23/08/2019	16/07/2021
Trend and Regime shifts	−7.973	−9.482	2	04/05/2007	28/08/2009	23/10/2015	08/12/2017	16/07/2021

Notes. The critical values for the Maki test depend on the number of regressors as well as the number of selected breaks. In performing the Maki test we selected a maximum number of breaks equal to 5 (that is m = 5), a stock market index as a dependent variable and the other four market indexes as independent variables (that is RV = 4): the critical values reported in this table correspond to these criteria and can be found also in Maki (2012).** denotes rejection of the null hypothesis of no cointegration at 5 % critical value.

The presence of structural breaks in the long-run relationship might have been caused by political change,[17] the behaviour of economic agents[18] and some shocks.[19] Furthermore, regarding the energy sector, the existence of a long-run relationship with structural break as detected in Panel D of Table 9 might be due to the fact that that major countries in South America have committed to reducing their carbon emissions and increasing the share of energy renewables as signatories of the Paris Agreement in 2015 (Peng et al., 2024). Major LAC companies operating in the energy sector may therefore have implemented actions to use new production technologies to reduce fossil fuel CO_2 emissions which might, in turn, have resulted in lower variable costs for all energy companies resulting in the same economic performance. Secondly, before starting the transition towards green technologies, large LAC companies operating in the energy sector have used the same commodities (such as coal products, natural gas and crude oil) to satisfy the demand of energy from households and firms. As the prices of these commodities are determined in the international markets, these companies might have encountered similar variable costs in their production, meaning the equity market performances of these firms might have behaved quite similarly over the long run. Panel E of Table 9 reports evidence of cointegration among LAC financial equity indices when the model with regime shift model is used. Regime shifts are fundamental changes in the market structure or macroeconomic environment: therefore, the instability of the long-run relationship among these sectoral financial equity indices might be due to their sensitivity to these fundamental changes.

7.2. An application of principal component analysis (PCA)

The intermittent evidence of long-run relationships revealed via the dynamic cointegration analysis, as well as the general absence of cointegration as revealed with the static cointegration approach, led us to examine market co-movements through the application of Principal Component Analysis (PCA). Not only did this enable us to check the robustness of the dynamic cointegration analysis, the additional goal of the PCA is also to identify factors (i.e. principal components) that can explain the variability of the entire set of variables. The application of the PCA can be summarised using several steps.[20] Firstly, principal components are generated as a linear combination of a set of variables. Secondly, these principal components are ordered with the first principal component (i.e. PC1) explaining most of the variance within the data set of variables and the last principal component the least variance. As pointed out by Gilmore et al. (2008), the larger the proportion of the variance explained by the PC1, the more evidence there is about equity markets' co-movements.

The proportion of the total variation in the equity indices returns explained by principal components was used as a measure of co-movements among LAC equity markets.[21] We implemented the PCA using both a dynamic as well as a static modality in the case of LAC national and sectoral equity indices. In the case of the former, the results of the dynamic analysis were plotted to provide a graphical illustration of the time path of co-movements. In this dynamic modality, we ran the PCA using two alternative dynamic approaches – the recursive and the rolling window. In the context of the recursive approach, the PCA was performed with an initial period of three years of data to calculate the principal components before extending this period by an additional year to calculate the principal components again. This exercise of extending the initial period recursively was repeated until the end of data was reached. In a rolling window context, we performed the PCA by using windows with a starting period of three years and then rolling that period ahead by adding one more year and removing the first year.

The findings of the dynamic recursive approach are shown in Panel A of Fig. 8. Initially, PC1 explained 79.12 % of the total variation in the set of LAC equity market indices, which then started to gradually decline, despite a new peak in 2011 (76.78 %), throughout the subsequent period.[22] With the rolling window approach, we performed the PCA for a total of sixteen rolling windows[23] and present the results in Panel B of Fig. 8. The PC1 in Panel B of Fig. 8 shows that the impact of the 2007–2009 global financial crisis might have contributed to the increase in the total variation of the LAC equity market returns from 79.02 % in 2007 to 94.02 % in 2008: during that period there was a substantial degree of market co-movement among the five LAC national equity markets. During the period following the end of the 2007–2009 global financial crisis the significance of the PC1 in explaining the total variation of the LAC equity market returns was quite volatile: total variation of returns reached another peak in the second half of 2011 (94.02 %) but also

[17] The political systems of the LAC in the first decade of the 2000s were mainly characterised by the prevalence of centre-left parties that formed governments that benefited from high commodity prices and implemented social and economic reforms. The second part of the 2010s witnessed the political sentiment that resulted in successive conservative governments.

[18] Since the beginning of the 2000s, LAC have seen a reduction of people living in poverty by almost a half, and a general increase in the size of the middle class, despite some variation across countries. The deceleration of growth during 2014–2019 following the decline of commodity prices, as well as the fall in economic activity caused by the Covid-19 crisis, impacted the behaviour of households and firms across LAC.

[19] The commodity boom of the first decade of the 2000s was the major factor that contributed to wide economic growth of the LAC since the end of the 1997 Asian financial crisis. As LAC are important commodity exporters, they greatly benefited through the increases in export revenues, resulting in strong economic growth (Campos, 2019).

[20] See Appendix B for a detailed discussion of the PCA methodology.

[21] Several empirical studies have used PCA to analyse co-movement patterns of equity market returns by focusing on emerging markets in Latin America (see, for instance, Meric, Leal, Ratner, & Meric, 2001) as well as Eastern Europe (see, for instance, Gilmore et al., 2008).

[22] All the principal components but the PC1 were found to have eigenvalues less than unity, therefore just only the PC1 was relevant for the PCA performed by using the recursive approach.

[23] For each rolling window, the eigenvalue of the PC1 was the one with a value greater than unity, whereas all the other principal components were found to have eigenvalues lower than 1. Therefore, also in the case of the rolling window approach, just only the PC1 was relevant in terms of interpreting the results of the PCA.

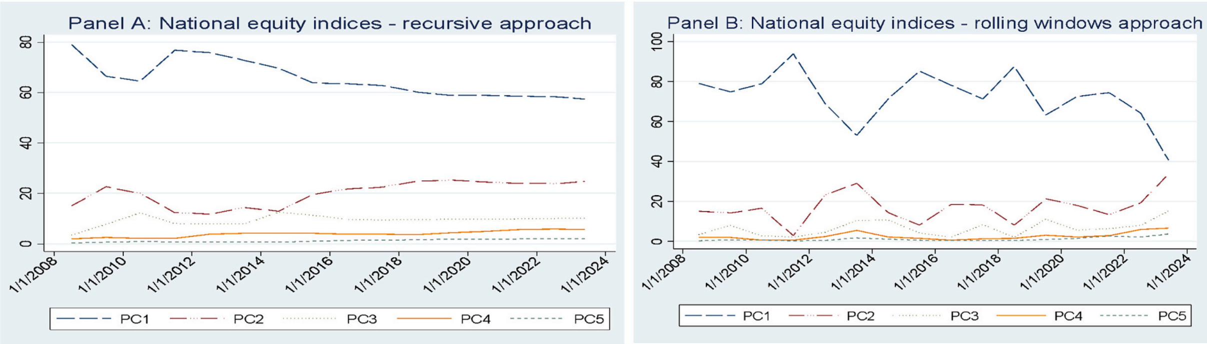

Fig. 8. Dynamic principal component analysis: national equity indices.

experienced the lowest value ever (53.09 %) in the second half of 2013, although a reverting trend seems to take place from the end of 2013. Finally, when comparing the findings of the dynamic cointegration analysis among the national equity markets of LAC with the dynamic PCA, we could say that there is some consistency in terms of findings when a rolling window approach is used in both cointegration and PCA. In other words, the findings of the rolling window cointegration (see Fig. 6) highlighted evidence of episodic long-run relationships whereas the rolling windows PCA highlighted that the PC1 accounted for a substantial degree of market linkages (Panel B of Fig. 8). On the other hand, the findings of the recursive dynamic cointegration approach (Fig. 4) are less consistent with the results of the recursive dynamic PCA (Fig. 8 – Panel A). While the former found evidence of long-run relationships among the LAC national equity market indices in the first part of the period of analysis only, the latter demonstrates that the CP1 accounts for a consistent proportion of the variance in the set of market returns. This variance is usually interpreted as an indication of a substantial degree of market linkages (Gilmore et al., 2008).

Figs. 9 to 13 report the results of the dynamic PCA performed at sectoral level. Based on these figures we can draw the following conclusions. Starting with the recursive approach of the PCA, Panels A of Figs. 9 to 13 show that first principal component explains larger variation in the set of sectoral indices returns in the first part of the period of analysis. In addition to this, Panels A of Figs. 9 to 13 also show that basic materials and financials are the sectoral indices where the total variation explained by the PC1 is usually larger than the consumer cyclicals, energy and industrials: therefore, the degree of market linkages in the former two sectors appear to be stronger in comparison to market linkages in the latter three sectors. The findings of the recursive approach, as detected by the dynamic cointegration analysis, are therefore partially consistent with the PCA recursive approach: the former did reveal intermittent co-movements among the LAC financial equity indices as well as among the LAC industrial equity indices and among the energy indices. However, the PAC analysis, in its recursive approach, also revealed that the degree of market linkages was higher in the financial indices in comparison to other LAC sectors.

Moving to the findings of the rolling windows approach of the dynamic PCA, Panels B of Figs. 9 to 13 show two relevant, time-evolving patterns of the market linkages at sectoral level. Firstly, in the first period of analysis, we observe that the PC1 explains the high levels in the total variability of returns in the case of basic materials, financials as well as industrials sectoral equity indices for a prolonged period. Values of the PC1 for consumer cyclicals and energy are generally lower and more volatile: the first three groups of sectoral indices therefore seem to have a substantial degree of market linkages in comparison to the ones related to the former two groups. Secondly, it is noteworthy that, for all sectoral indices, the PC1 values show a declining trend that, in most of the cases, coincides with the end of the 2007–2009 global financial crisis. The findings of the rolling windows PCA are partially consistent with the results of the rolling windows cointegration analysis as presented in Panels A and B of Fig. 7: both Panels, irrespective of the normalised cointegration statistics used, show that the basic materials, energy and financials of the LAC sectoral indices are characterised by

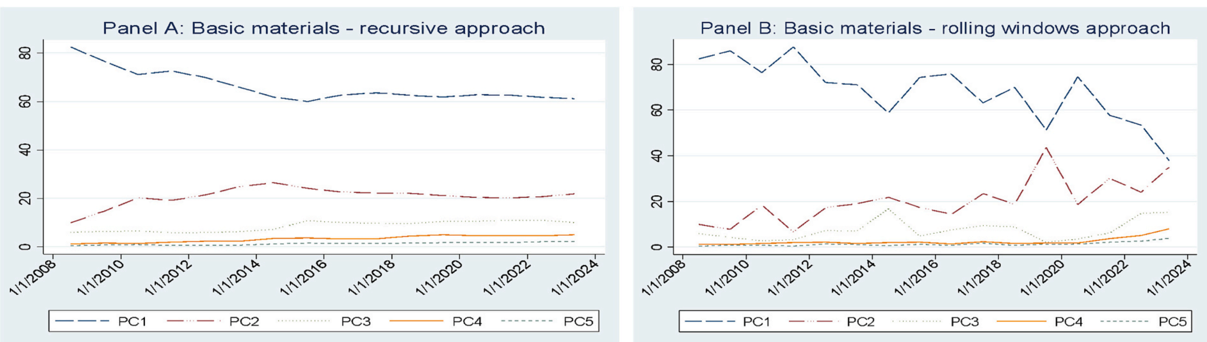

Fig. 9. Dynamic principal component analysis: basic materials equity indices.

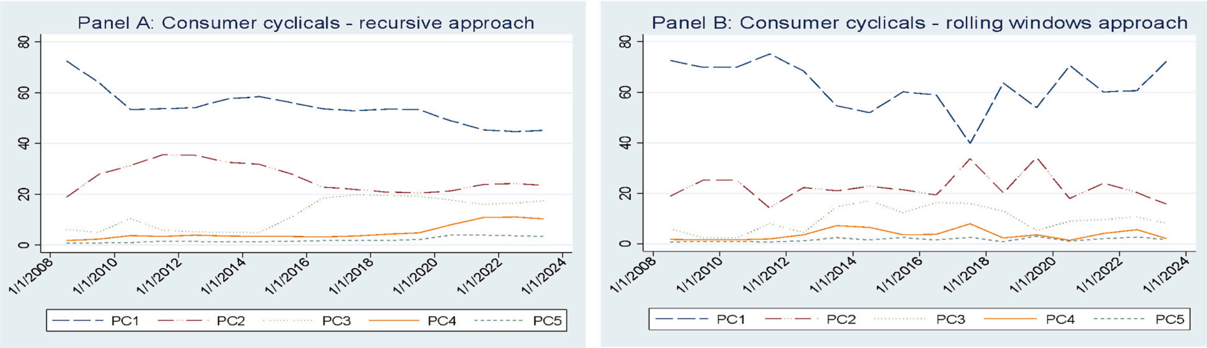

Fig. 10. Dynamic principal component analysis: consumer cyclicals equity indices.

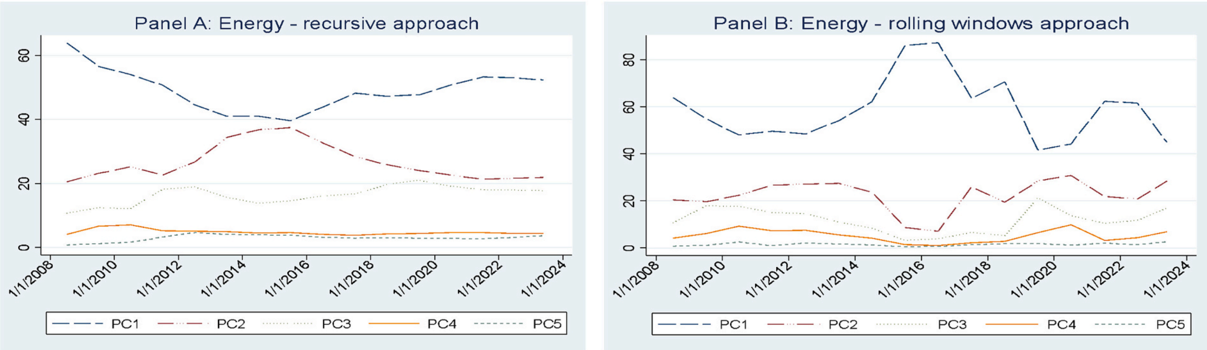

Fig. 11. Dynamic principal component analysis: energy equity indices.

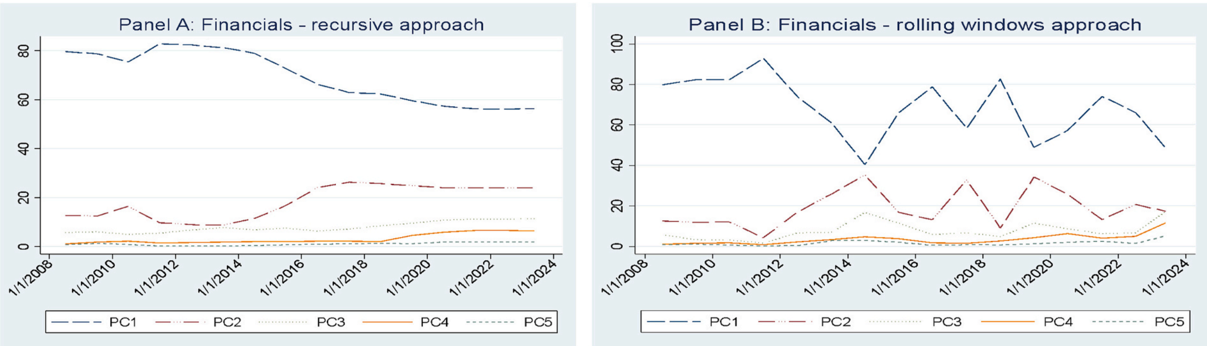

Fig. 12. Dynamic principal component analysis: financials equity indices.

intermittent cointegration in the first period of analysis. This thus confirms the results of the PCA for both groups of indices in terms of substantial market co-movements.

To complement the results of the dynamic application of the PCA, we applied the same methodology in a static mode. The results are presented in Table 10 which shows that the first principal component (PC1) and the second principal component (PC2) are relevant for our analysis as the associated eigenvalues are greater than 1. Eigenvalues values for all other remaining principal components are less than unity and, following Kaiser's significance rule, only principal components greater than unity are taken for analysis. In accordance with Kaiser's rule, the first principal component (Column 1 of Table 10) is statistically significant: in particular, the Chilean and Peruvian equity market returns have the highest factor loadings, accounting for 57.4 % of the total variation in this set of the LAC equity market returns. On the other hand, the PC2 is also statistically significant and explains 24.6 % of the total variation in this set of the LAC equity market returns.

In Tables 11 to 15, we present the results of the PCA as a static mode of analysis applied to the LAC equity indices aggregated by sector. First, we observe that in each table only the first two eigenvalues of the first and second principal components are bigger than

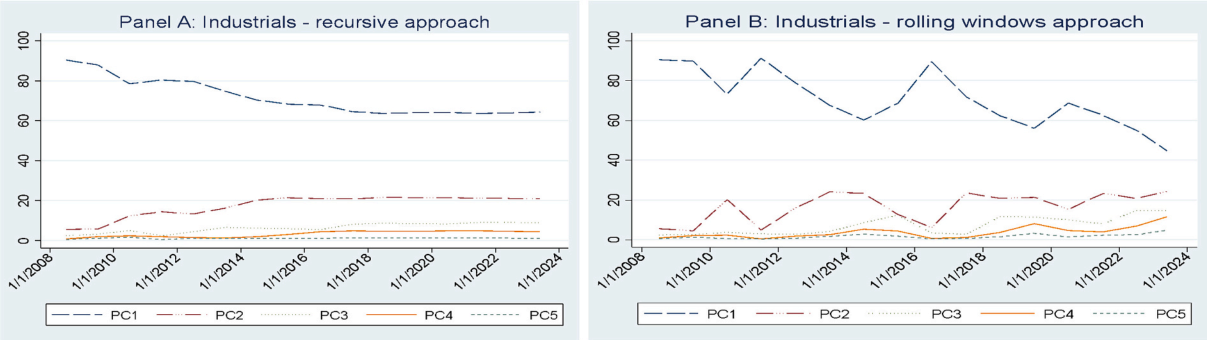

Fig. 13. Dynamic principal component analysis: industrials equity indices.

Table 10
Principal components analysis - national equity market indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	0.062	0.868	0.110	0.402	0.260
Brazil	0.478	−0.294	−0.426	0.707	−0.046
Chile	0.534	0.291	0.074	−0.245	−0.751
Colombia	0.447	−0.258	0.807	0.094	0.266
Peru	0.530	0.088	−0.384	−0.518	0.542
Eigenvalue	2.869	1.232	0.510	0.285	0.101
Variance explained	0.574	0.246	0.102	0.057	0.02

Notes: The eigenvalue was 2.869 for the first principal component, 1.232 for the second principal component, 0.510 for the third principal component, 0.285 for the fourth principal component and 0.101 for the fifth principal component. Applying Kaiser's rule, we retain only the principal components whose eigenvalues exceed 1.

unity: therefore, the PC1 and PC2 are the only principal components statistically significant in explaining the total variation in the LAC sectoral equity indices returns. We observe that the first principal component explains the largest variation in the set of sectoral returns related to the basic materials (Table 11) equity indices (61.1 %) as well as the industrial equity (Table 15) indices (64.3 %). On the other hand, the total variation in the financials (Table 14), energy (Table 13) and consumer cyclicals (Table 12) equity returns that can be explained by the PC1 are much lower with values of 56.3 %, 52.3 % and 36.5 % respectively.

8. Conclusions

In this study we examined long-run relationships among the major LAC equity markets. We used weekly data for MERVAL (Argentina), IBOVESPA (Brazil), COLCAP (Colombia), IGPA (Chile) and IGBVL (Peru) indices over the 2005 to 2023 period. In addition, and quite differently to most of the studies in this area, we also carried out sectoral analysis for the basic materials, consumer cyclicals, energy, financials and industrial equity indices of each country.

To allow for some dynamics in the linkages among LAC equity markets, we applied the Hansen and Johansen (1992) recursive cointegration method and its rolling windows variation developed by Pascual (2003). To check for the robustness of the results both at national and sectoral level, we also applied the standard static cointegration methodology (Johansen, 1988; Johansen & Juselius, 1990). We further complemented this analysis with two tests to ascertain the presence of structural breaks in the long-run relationships of LAC equity markets both at national and sectoral level: the Gregory and Hansen (1996) and Maki (2012) cointegration tests with structural breaks. Finally, the robustness of the results was also tested by using the principal component analysis (PCA) technique.

The application of the dynamic cointegration analysis based either on the recursive or rolling window approach indicates that the LAC equity markets witnessed some intermittent periods of cointegration. For example, from mid-2011 to mid-2014, the cointegration vectors are statistically significant and therefore there is evidence of temporary long-run relationships across the equity market indices that eliminate the scope for equity portfolio diversification across individual LAC. This period coincides with some economic events that affected the LAC economies such as the end of the commodity boom that fuelled economic growth in Latin America since the first half of 2000s; the effects of the MILA agreement in 2009 that integrated their stock exchanges; as well as relevant political changes across many LAC following the end of the 2007–2009 global financial crisis. We also found some evidence of intermittent periods of cointegration at the sectoral level across the equity indices of the LAC. To further check the robustness of the dynamic cointegration analysis, we implemented a dynamic as well as a static principal component analysis which demonstrated that the first principal component was able to explain a large proportion of the variation of the LACs' national, as well as sectoral, equity indices. On the other hand, the findings of the static cointegration analysis suggest that, over our period of investigation, the LAC equity markets as a whole

Table 11
Principal components analysis - basic materials equity indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	−0.020	0.919	0.387	0.061	0.002
Brazil	0.486	0.177	−0.054	0.078	0.129
Chile	0.533	0.08	−0.074	−0.576	0.609
Colombia	0.426	−0.335	0.765	0.335	0.082
Peru	0.544	0.06	−0.06	−0.302	−0.777
Eigenvalue	3.057	1.091	0.499	0.244	0.107
Variance explained	0.611	0.218	0.099	0.048	0.021

Notes: The eigenvalue was 2.676 for the first principal component, 0.777 for the second principal component, 0.607 for the third principal component, 0.492 for the fourth principal component and 0.445 for the fifth principal component. Applying Kaiser's rule, for our analysis we retain only the principal components whose eigenvalues exceed 1.

Table 12
Principal components analysis - consumer cyclicals equity indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	0.411	0.919	−0.533	0.723	0.150
Brazil	0.609	0.177	−0.069	−0.239	−0.748
Chile	0.570	0.08	−0.06	−0.498	0.641
Colombia	0.318	−0.335	0.816	0.413	0.077
Peru	0.178	0.06	0.202	0.008	0.018
Eigenvalue	1.826	0.974	0.934	0.790	0.473
Variance explained	0.365	0.194	0.186	0.158	0.094

Notes: The eigenvalue was 1.826 for the first principal component, 0.974 for the second principal component, 0.934 for the third principal component, 0.790 for the fourth principal component and 0.473for the fifth principal component. Applying Kaiser's rule, for our analysis we retain only the principal components whose eigenvalues exceed 1.

Table 13
Principal components analysis - energy equity indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	0.213	0.799	0.412	0.314	−0.216
Brazil	0.548	−0.271	0.138	−0.344	−0.698
Chile	0.505	−0.165	−0.469	0.705	0.020
Colombia	0.462	0.412	−0.432	−0.533	0.566
Peru	0.430	−0.300	0.635	0.028	0.057
Eigenvalue	2.616	1.096	0.886	0.216	0.184
Variance explained	0.523	0.219	0.177	0.043	0.036

Notes: The eigenvalue was 1.231 for the first principal component, 0.940 for the second principal component, 0.789 for the third principal component, 0.521 for the fourth principal component and 0.429 for the fifth principal component. Applying Kaiser's rule, we retain only the principal components whose eigenvalues exceed 1.

do not exhibit any evidence of long-run relationships. This clearly indicates that there is some potential for international portfolio equity diversification benefits across our LAC markets. Similarly, our sectoral analysis points to specific diversification opportunities across most of the sectors.

Some studies (see, for instance, Gilmore & McManus, 2003) have pointed out that lack of cointegration among equity markets might be a consequence of weak economic integration among the respective economies. Considering those findings, we could assume that even if the LAC have made substantial progress in co-ordinating their economies with the Mercosur, the current state of their economic integration is still far from producing the effect of bringing their equity markets into a long-run relationship. Finally, it seems that there could be some benefits of international portfolio diversification in LAC equity markets. To make these benefits quantifiable, specific further work is needed by applying portfolio diversification strategies.

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Table 14
Principal components analysis - financials equity indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	0.329	0.653	−0.380	0.554	0.107
Brazil	0.411	−0.495	0.435	0.608	0.159
Chile	0.571	−0.061	−0.081	−0.169	−0.796
Colombia	0.471	−0.347	−0.545	−0.340	0.493
Peru	0.415	0.450	0.601	−0.421	0.292
Eigenvalue	2.814	1.195	0.566	0.324	0.098
Variance explained	0.563	0.239	0.113	0.065	0.019

Notes: The eigenvalue was 2.696 for the first principal component, 0.773 for the second principal component, 0.715 for the third principal component, 0.434 for the fourth principal component and 0.380 for the fifth principal component. Applying Kaiser's rule, we retain only the principal components whose eigenvalues exceed 1.

Table 15
Principal components analysis - industrial equity indices.

Equity market	First principal component	Second principal component	Third principal component	Fourth principal component	Fifth principal component
Argentina	0.057	0.962	0.091	0.250	−0.061
Brazil	0.463	−0.004	0.787	−0.366	0.173
Chile	0.474	0.191	−0.563	−0.644	−0.065
Colombia	0.523	−0.124	−0.228	0.481	0.652
Peru	0.531	−0.150	0.032	0.393	−0.734
Eigenvalue	3.219	1.049	0.446	0.225	0.059
Variance explained	0.643	0.209	0.089	0.045	0.011

Notes: The eigenvalue was 2.334 for the first principal component, 0.877 for the second principal component, 0.768 for the third principal component, 0.612 for the fourth principal component and 0.407 for the fifth principal component. Applying Kaiser's rule, we retain only the principal components whose eigenvalues exceed 1.

In our research we did not make use of generative AI or AI-assisted technologies such as ChatGPT or similar services.

CRediT authorship contribution statement

Francesco Guidi: Writing – review & editing, Writing – original draft, Validation, Supervision, Software, Project administration, Methodology, Formal analysis, Data curation. **Giuseppina Madonia:** Writing – review & editing, Writing – original draft, Validation, Supervision, Methodology, Investigation, Formal analysis, Conceptualization. **Sohan Sarwar:** Writing – review & editing, Visualization, Supervision, Software, Methodology, Investigation, Data curation, Conceptualization.

Declaration of competing interest

We have nothing to declare.

Appendix A

The standard model used to detect the presence of cointegration with no structural break is represented as follows

$$y_{1t} = \mu + \alpha^T y_{2t} + e_t \tag{A1}$$

Eq. (A1) assumes that the parameters μ and α are time invariant. The presence of cointegration with structural breaks is investigated by testing for changes in the intercept μ and/or changes to the slope α (Gregory & Hansen, 1996). This test does require the modification of the standard cointegration model (Eq. (A1)) by adding a dummy variable $\varphi_{t\tau}$ which can take either 1 or 0, that is:

$$\varphi_{t\tau} = \{0 \text{ } if \text{ } t \leq [n \text{ } \tau] \text{ } 1 \text{ } if \text{ } t > [n \text{ } \tau] \tag{A2}$$

where the timing of the change point is the unknown parameter $\iota \in (0 \text{ } 1)$. Alternative models can be used to detect the presence of structural change in the cointegration relationship (Gregory & Hansen, 1996). The simple case (i.e. the *level shift* model expressed as *model C*) assumes that, under the alternative hypothesis of cointegration with structural break, there is a change in the intercept μ, while the slope coefficient α is held time invariant. The mathematical notation of m*odel C* is as follows:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^T y_{2t} + e_t \tag{A3}$$

where μ_1 represents the intercept before the shift, and μ_2 represents the change in the intercept at the time of the shift (Gregory & Hansen, 1996).

The *level shift with trend (*expressed as *model C/T*) is an additional alternative model which departs from *model C* as the former includes a time trend parameter *t* (Gregory & Hansen, 1996). Therefore, the *model C/T* can be represented as follows:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t \tag{A4}$$

Lastly, an additional alternative model is the one with a regime shift where this structural change allows the slope vector to shift as well. This is called *regime shift* model (expressed as *model C/S*) and it is represented as follows:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t}\varphi_{t\tau} + e_t \tag{A5}$$

where μ_1 and μ_2 are as in *models C* and *model C/T*, whist α_1 denotes the cointegrating slope coefficients before the regime shift and α_2 denotes the change in the slope coefficients.

In each of these models, we test the null hypothesis of no cointegration versus the alternative of cointegration by taking into account a possible regime shift in the LAC equity markets. A rejection of the null hypothesis implies that there are long-run relationships among these equity markets.

Appendix B

The PCA is a technique applied to a set of variables to discover which variables form coherent subsets which are relatively independent of one another. The PCA of a set of variables $X_1, X_2, \ldots, X_m$ that generates a set of new variables; that is, the principal components $Y_1, Y_1, \ldots, Y_m$ with each component being a linear combination of the original variables, that is

$$Y_1 = b_{1,1}X_1 + b_{1,2}X_2 + \ldots + b_{1,m}X_m = b_1'X$$

$$Y_2 = b_{2,1}X_1 + b_{2,2}X_2 + \ldots + b_{2,m}X_m = b_2'X$$

$$\vdots$$

$$Y_m = b_{m,1}X_1 + b_{m,2}X_2 + \ldots + b_{m,m}X_m = b_m'X$$

The coefficients for the first principal component Y_1 are chosen as to make the variance of Y_1 as large as possible (Harris, 1975). In general, the coefficients for each of the principal components are chosen to maximise the variance of each principal component subject to the restriction that it be uncorrelated with scores on Y_1 trough Y_{i-1}, therefore:

$$Var(Y_i) = b_i'S_xb_i \; i = 1, 2, \ldots, m$$

$$Cov\,(Y_i, Y_k) = b_i'S_xb_k \; i, k = 1, 2, \ldots, m$$

where S_x is the covariance matrix of $X_1, X_2, \ldots, X_m$, Using the method of the Lagrangian multiplier, the coefficients b_i of each principal components Y_i can be determined by solving the following equations:

$$L_1 = b_1'S_xb_1 - \lambda\left(b_1'b_1 - 1\right)$$

$$L_2 = b_2'S_xb_2 - \lambda\left(b_2'b_2 - 1\right)$$

$$\vdots$$

$$L_m = b_m'S_xb_m - \lambda\left(b_m'b_m - 1\right)$$

This set of equations has a non-trivial solution if, and only if, the determinant $|S_x - \lambda I|b_i$ is equal to zero. Solving the Lagrangian equations by λ produces m roots: the roots with zero value indicating a linear dependence among the original variables. On the other hand, any one of the nonzero roots can be entered into the matrix equation $|S_x - \lambda_i I|b_i = 0$ and the resulting set of equations solved for the coefficient of b. Finally, since the first principal component Y_1 has been set out as the one with the larger variance, the coefficients of the first principal component will be the characteristic vector associated with the largest characteristic root of the characteristic equation λ_i. The second principal component Y_2 is calculated via the characteristic vector corresponding to the second largest characteristic root and, as it turns out, the same approach can be used to calculate all the remaining principal components (Harris, 1975). The proportion of the total variance explained by the *kth* principal component is then calculated as $\lambda_k/(\lambda_1 + \lambda_2 + \ldots + \lambda_m)$, where $k = 1, 2, \ldots, m$.

Data availability

Data will be made available on request.

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