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NEW EVIDENCE ON LONG-RUN OUTPUT CONVERGENCE AMONG LATIN AMERICAN COUNTRIES

Mark J. Holmes*
Waikato University Management School

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This study assesses long-run real per capita output convergence among selected Latin American countries. The empirical investigation, however, is based on an alternative approach. Strong convergence is determined on the basis of the first largest principal component, based on income differences with respect to a chosen base country, being stationary. The qualitative outcome of the test is invariant to the choice of base country and, compared to alternative multivariate tests for long-run convergence, this methodology places less demands on limited data sets. Using annual data for the period 1960-2000, strong convergence is confirmed for the Central American Common Market. However, an amended version of the test confirms weaker long-run convergence in the case of the Latin American Integration Association countries.

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I. Introduction

In recent years, economists have keenly debated the issue of whether or not per capita incomes across countries are converging. The neoclassical growth model predicts that countries will converge towards their balanced growth paths where per capita growth is inversely related to the starting level of income per capita. Early studies by Barro (1991), Barro and Sala-i-Martin (1991,1992), Baumol (1986), Sala-i-Martin (1996) and others that consider convergence across countries, US states and European regions, argue that in most instances the rate annual rate of convergence is roughly 2%. This is confirmed by studies such as Mankiw et al.

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* Department of Economics, Waikato University Management School, Hamilton, New Zealand. Email: holmesmj@mngt.waikato.ac.nz. I would like to express my gratitude to the co-editor Germán Coloma and three anonymous referees for their very helpful comments. Any remaining errors are my own.

(1992) who investigate conditional convergence that allows for population growth and capital accumulation. More recent studies have offered mixed evidence on this question. For instance, Quah (1996) questions the 2% convergence rate and argues that convergence will take place within relatively homogenous convergence clubs. McCoskey (2002) suggests that convergence clubs and regional homogeneity is probably unresolved with respect to less developed countries (LDCs) where geographic proximity and cross-national economic interdependence will cause groups of LDCs to grow or falter as one. As noted by Dobson and Ramlogan (2002), little is known about the convergence process among LDCs and the limited range of studies that have considered LDCs have proceeded at a highly aggregated level (Khan and Kumar 1993) or have focused on convergence within a particular country (Ferreira 2000, Nagaraj et al. 2000, Choi and Li 2000). The purpose of this paper is to examine convergence among Latin American countries where we assess the possibility of convergence clubs within LDCs based on common characteristics regarding international trade arrangements.

The recent study by Dobson and Ramlogan (2002) investigates convergence among Latin American countries over the study period 1960-90. They find evidence of unconditional beta convergence (poor countries growing faster than richer countries towards a common steady state) but not sigma convergence (distribution of income becoming more equal) across the full study period. However, by looking at sub-periods, they find that the rates of conditional convergence towards individual steady states are highest during the 1970s-mid 1980s. In addition to this, Dobson and Ramlogan conclude that the estimates of convergence may be sensitive to how GDP is measured.

The question of whether trade liberalization is associated with income convergence remains unresolved both in terms of theory and evidence.[1] Using annual data on real per capita GDP for a total sample of sixteen Latin American countries,[2] this study offers an empirical assessment of whether long-run income convergence among LDCs has been achieved by countries that have participated in the Latin American Integration Association (LAIA) and Central American Common Market (CACM). The LAIA was formed in 1980 by Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Paraguay, Peru, Uruguay, and Venezuela,

[1] See, for example, Slaughter (2001) and references therein.

[2] The full list of countries includes Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Paraguay, Peru, Uruguay and Venezuela.

taking over the duties of the Latin American Free Trade Association (LAFTA), which had been created in 1960 to establish a common market for its member nations through progressive tariff reductions until the elimination of tariff barriers by 1973. In 1969 the deadline was extended until 1980, at which time the plan was scrapped and the new organization, the LAIA, was created by the Treaty of Montevideo. It has the more limited goal of encouraging free trade, with no deadline for the institution of a common market. Economic hardship in Argentina, Brazil, and many other member nations has made LAIA's task difficult. The CACM was created in 1960 by a treaty between Guatemala, Honduras, Nicaragua, El Salvador, and later Costa Rica. By the mid-1960s the group had made advances toward economic integration, and by 1970 trade between member nations had increased more than tenfold over 1960 levels. During the same period, imports doubled and a common tariff was established for 98% of the trade with non-member countries. In 1967, it was decided that CACM, together with the Latin American Free Trade Association, would be the basis for a comprehensive Latin American common market. However, by the early 1990s little progress toward a Latin American common market had been made, in part because of internal and internecine strife, and in part because CACM economies were competitive, not complementary. Nonetheless, the CACM with a stronger focus on creating a common market than the LAIA, has been judged more successful at lowering trade barriers than other Latin American groupings.

While this study is motivated by the desire to throw more light on the issue of convergence among LDCs, there are further reasons of interest attached to this study. First, a key contribution is in terms of the methodology employed. The tests for income convergence are on the basis of whether the largest principal component, based on benchmark deviations from base country output, is stationary or not. This methodology, initially advocated by Snell (1996), offers a number of advantages over existing tests for convergence. Unlike the estimation of bivariate equations, the outcome of this test for convergence is not critically dependent on the choice of base country. Also, there are advantages over alternative common trends methods based on Johansen (1988) and Stock and Watson (1988), which can suffer from low test power on account of data limitations, as well as principal components analyses that search for integration using arbitrary methods to determine the 'significance' of given components. Second, the concept of convergence in the context of the groupings we analyze is important. Essentially, this study tests the hypothesis that convergence is a phenomenon where experience as trading partners or geographical location has the potential to bind economies together. Given that these agreements have sought to promote integration as part of their long-term objectives, the absence of convergence would justify the need for proactive policies to promote growth and reduce income inequalities. If one finds that the incomes of countries within these groups have converged, then it becomes more difficult to justify regional development policy in terms of economic efficiency (Dobson and Ramlogan 2002). Third, the concept of convergence employed in this study differs from that employed by Barro, Sala-i-Martin and others. These studies define convergence with respect to poorer countries growing faster than richer countries towards some (common or individual) steady state. The notion of convergence employed in this study is based on testing whether per capita outputs move together over time with no tendency to drift further apart in the long-run following a deviation from equilibrium.

The paper is organized as follows. The following section briefly considers the literature on trade liberalization and income convergence. The groupings of countries used in this study are then outlined. Section III discusses the data and econometric methodology. This leads to a new categorization of types of real convergence based on the stationarity of the first largest principal component. Section IV reports and discusses the results. The evidence suggests that long-run convergence is strongest among the CACM for whom the first largest principal component is stationary. On the other hand, the LAIA countries are weakly convergent. Section V concludes.

II. Trade, convergence and international agreements

The traditional approach to the convergence debate concerns poor countries catching up with rich ones. In the approaches taken by studies such as Barro (1991), Barro and Sala-i-Martin (1991,1992), Baumol (1986) and Sala-i-Martin (1996), a cross-section of growth rates are regressed on income levels and the estimated coefficient informs on the rate at which poor countries catch up with those richer. Quah (1996) argues that the conventional analyses miss key aspects of growth and convergence. Moreover, it is argued that the key issue is what happens to the cross-sectional distribution of economies, not whether an economy tends towards its own steady state. Quah therefore considers issues of persistence and stratification in the context of convergence clubs forming where the cross-section polarizes into twin peaks of rich and poor. The economic forces that drive this notion of convergence include factors such as capital market imperfections, country size, club formation, etc.

Structural and institutional factors are crucial in forming the background against which long-run linkages between countries can exist. As pointed out by Slaughter (2001), many papers on convergence cannot analyze the role of international trade because they assume a 'Solow world' in which countries produce a single aggregate good independently of each other. Moreover, convergence arises from capital stock convergence. However, trade theory that draws on and develops some of the arguments belonging to the factor price equalization theorem, Heckscher-Ohlin models, Stolper-Samuelson effects or Rybczynski theorem offers an ambiguous prediction as to whether or not trade liberalization will cause per capita incomes to converge or diverge. The convergence of factor prices via the factor price equilibrium theorem depends on cross-country tastes, technology and endowments. It is argued that trade liberalization has an ambiguous effect on endowments of labor and capital (see, for example, Findlay 1984). Trade liberalization may reduce investment risk particularly in poorer countries (see, for example, Lane 1997). Divergence may occur through the Stolper-Samuelson effects of liberalization on capital rentals where Baldwin (1992) argues that dynamic gains from trade will mean that richer countries that are well endowed with capital will experience increased capital rentals. Ventura (1997) argues that free trade may inhibit the onset of diminishing returns to investment where richer countries do not lose their incentive to invest as they would under autarky. Finally, income convergence will be affected by technology flows. Matsuyama (1996) argues that freer trade leads poorer countries to specialize in technologically-stagnant products because they lack the resources to engage in the production of high-technology products.

Empirical evidence on trade and income convergence is also mixed. Ben-David (1993, 1996) and Sachs and Warner (1995) find that international trade causes convergence. Sachs and Warner point to the convergence club of economies linked by international trade. Ben-David (1996) finds that it is the wealthier countries that trade significantly who are characterized by per capita convergence. Ben-David (1993) analyses five episodes of post-1945 trade liberalization and finds that income convergence generally shrank after liberalization started. On the other hand, Bernard and Jones (1996) find that freer trade causes incomes to diverge while Slaughter (2001), using a sample of developed countries and LDCs, finds no strong, systematic link between trade liberalization and convergence. Indeed, Slaughter suggests that much of the evidence indicates that trade liberalization diverges incomes among the liberalizers.

III. Data and methodology

This study employs data for annual per capita real GDP (US$) for each of the sample of countries for study periods of 1960 up to 2000. All data are obtained from the Penn World Tables version 6.1. The following study periods are considered: the full study period of 1960-2000, 1981-2000 which represents the period of operation for the LAIA[3], and 1960-1980 to analyse convergence among the LAIA countries before their agreement became operational. As well as examining groupings based on the CACM and LAIA countries, this study also considers the full sample of countries taken together (*All*) as well as groupings based on geographical location within Latin America namely, Colombia, Costa Rica, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua and Venezuela (*North*) along with Argentina, Bolivia, Brazil, Chile, Paraguay, Peru and Uruguay (*South*). The exclusion of certain countries from some of the groups is driven by data availability over the full study period. Using this data, this study employs a two-stage testing procedure for income convergence. The first stage draws on a technique, developed by Snell (1996), which is an extension of the principal components methodology, based on testing for the stationarity of the first largest principal component (LPC) of benchmark deviations from base country output for each group in turn. This test can confirm long-run convergence where all series move in tandem over the long-run. This can be described as *strong convergence*. The second stage applies if stage one finds against strong convergence. Principal components are computed for each group where per capita incomes are expressed in levels rather than differences from base country and the number of common shared trends are calculated. This second stage searches for evidence of a single common shared trend driving the output series. This would confirm weak convergence because, unlike stage one, homogeneity between the countries has been relaxed.

With regard to the first stage of the convergence test, suppose $n + 1$ countries constitute the sample of a given group. The benchmark deviations are defined as

$$(y_i - y_G)_t = u_{it}, \tag{1}$$

where y_{it} and y_{Gt} respectively denote the natural logarithm of the real per capita income of country i and the chosen base-country, and $i = 1, 2,\dots, n$. Let X_t be an

[3] The LAIA study period covers 1981-2000 because the 1980 Treaty of Montevideo became operational in 1981.

(nx1) vector of random variables, namely the u_{it}'s for each of the n countries, which may be integrated up to order one. The principal components technique addresses the question of how much interdependence there is in the n variables contained in X_t. We can construct n linearly independent principal components which collectively explain all of the variation in X_t where each component is itself a linear combination of the u_{it}'s.[4] Since I(1) variables have infinite variances, whereas stationary, I(0), variables have constant variances, it follows that the first LPC, which explains the largest share of the variation in X_t, is the most likely to be I(1) and so corresponds to the notion of a common trend (Stock and Watson 1988). However, if the first LPC is I(0) then all the remaining principal components will also be stationary and there are no common trends which suggests that the u_{it}'s contained in X_t are themselves stationary. This will confirm real convergence with the base-country across the sample of n benchmark deviations.

More formally, following Stock and Watson (1988) we can argue that each element of X_t may be written as a linear combination of $k \leq n$ independent common trends which are I(1), and $(n-k)$ stationary components which correspond to the set of $(n-k)$ cointegrating vectors among the u_{it}'s. The k vector of common trends and $(n-k)$x1 vector of stationary components may respectively be written as

$$\tau_t = \alpha' X_t, \tag{2}$$

$$\xi_t = \beta' X_t, \tag{3}$$

where α is an $(n-k)$ matrix of full column rank, β is an nx$(n-k)$ matrix that forms the $(n-k)$ cointegrating vectors, $\alpha'\alpha = I$ and $\alpha'\beta = 0$. If there are k common trends, it can be shown that the k LPCs of X_t may be written as

$$\tau_t^* = X_t^{*\prime} \alpha^*, \tag{4}$$

where X_t^* is a vector of observations on the u_{it}'s in mean deviation form, α^* represents the k eigenvectors corresponding to the largest eigenvalues of X_t and is defined as αR where R is an arbitrary, orthogonal (kxk) matrix of full rank. This relationship guarantees that under the null hypothesis of k common trends, each of the k LPC's will be I(1). Similarly, for the $(n-k)$ remaining principal components, it can be shown that

[4] See, for example, Child (1970).

$$\xi_t^* = X_t^{*'} \beta^*, \tag{5}$$

where β^* corresponds to the $(n - k)$ eigenvectors that provide the $(n - k)$ smallest principal components and is defined as βS where S is an arbitrary orthogonal $(n - k)$x$(n$-$k)$ matrix.

The first LPC will be I(1) provided there is at least one common trend among the u_{it}'s contained in X_t. We can therefore test the null hypothesis that the first LPC is non-stationary against the alternative hypothesis that the first LPC is I(0). Rejection of the null means that all principal components are stationary and so there are no common trends among the u_{it}'s contained in X_t. This confirms convergence with respect to the base-country across the sample. To test the stationarity of the first LPC we can use the familiar Augmented Dickey-Fuller (ADF) test based on

$$\Delta z_{1t} = \rho z_{1t-1} + \sum_{i=1}^{p} \gamma \Delta z_{1t-i} + e_t, \tag{6}$$

where the first LPC is calculated as $z_1 = \alpha_1^{*} X_t^{*}$ using α_1^* as the first column of α^*, and e_t is a white noise error term. If we find that z_1 is trend stationary only, this will not confirm convergence because for at least one series in the sample, the difference from base country is growing over time. This would imply the presence of at least two common shared trends among the X_t ‘s.

This notion of convergence can be seen in the context of the Bernard and Durlauf (1995) definition of convergence in a stochastic environment where the long-run forecasts of the benchmark deviations tend to zero as the forecast horizon tends to infinity. If each y is I(1) or first difference stationary, then convergence implies that each $(y_i - y_G)_t = u_{it}$ is a stationary process (since the per capita income series are indices having different bases, we may allow the benchmark deviations to have different means) where each y_{it} and y_{Gt} is cointegrated with a cointegrating vector [1, -1].

An alternative way forward is to test for a single common trend among a series of I(1) variables $(y_{1t}, y_{2t}, ..., y_{nt}, y_{Gt})$ where convergence is confirmed through the presence of n cointegrating vectors among the $n + 1$ countries. The advantage of examining the stationarity of the first LPC is that, unlike the Johansen (1988) maximum likelihood procedure (and the Stock and Watson 1988 common trend framework), it does not require the estimation of a complete vector autoregression system (VAR).[5] The size and power of this test is not affected by the VAR being

[5] See, for example, Mills and Holmes (1999) who employ these methods to examine common trends among European output series during the Bretton Woods and Exchange Rate Mechanism eras.

constrained to an unreasonably low order on account of data limitations. This method also avoids the need for an entire sequence of tests for the stationarity of a multivariate system. As indicated by Snell, even if each test in the sequence had a reasonable chance of rejecting the false null, the procedure as a whole is likely to have low power. Another important issue is whether the choice of base country affects the outcome of the test. This methodology employed in this paper is based on a multivariate test for convergence that is not critically dependent upon the choice of base country. In one scenario we may find that the first LPC constructed from the n income differentials is stationary thereby suggesting that all $n + 1$ countries in the sample share the same common stochastic trend. It will not matter which country is used as base because the first LPC will still be stationary. If the first LPC is non-stationary, then there are at least two common stochastic trends among the sample of $n + 1$ countries with a maximum of n countries sharing the same trend. In this case, it is impossible to change to base country so that the first LPC is stationary.

The second stage of the test is applied if one is unable to reject the null that the first LPC based on differences with respect to base country is non-stationary. So far, under the LPC test, income differentials are constructed among the lines of $(y_i - \beta_i y_G)_t$ where $\beta_i = 1 \,\forall i$. Since the differentials are computed as $(y_i - y_G)_t$, this means that homogeneity has been imposed, i.e., $\beta_i = 1 \,\forall i$, before the test is conducted. *Strong convergence*, which is based on homogeneity, is therefore confirmed if the first LPC is stationary. Non-stationarity of the first LPC may occur because $\beta_i \neq 1$ in at least one case. Even if y_i and y_G are cointegrated, it may now be more appropriate to think of the long-run cointegrating relationship not being written as $y_{it} = y_{Gt} + u_{it}$ but rather as $y_{it} = \beta_i y_{Gt} + u_{it}$ instead. In the latter case, homogeneity has not been imposed and it is *weak convergence* that is being tested for where the variables used to construct the principal components are $y_1, \ldots\ldots\ldots, y_n, y_G$ instead of $(y_i - y_G) = u_i$ for $i = 1, \ldots.., n$. Moreover, it is possible that the y series are driven by a single common shared trend but without the homogeneity that strong long-run convergence implies.

To address the possibility of weak convergence, principal components are computed for the $n + 1$ countries expressed in levels rather than differences with respect to a base country. If the first and second LPCs are respectively non-stationary and stationary, this will suggest there are n cointegrating vectors present and therefore one $((n + 1) - n = 1)$ common shared trend. We can describe this as *weak convergence* because the first stage of the test described previously did not support convergence based on homogeneity, yet the second stage of the test found that the countries are nonetheless sharing the same long-run trend.[6] We may find that the third LPC is the first principal component that is stationary. In this case, we have $n-1$ cointegrating vectors present and this implies the presence of two $((n+1)-(n-1)=2)$ common trends among the $n+1$ countries. This is yet weaker evidence of convergence. In the extreme, we may find that none of the principal components are stationary. This implies that there are no cointegrating vectors and therefore $n+1$ common trends and the sample of $n+1$ countries. This would be consistent with zero long-run convergence or complete divergence.

Before proceeding to the results discussion, it is important to highlight some caveats associated with this methodology. The advantages over existing methods of testing for long-run convergence have been discussed, however the downside of this methodology concerns a standard criticism of principal component estimation and indeed of common stochastic trends. They are linear combinations of economic variables and so the economic interpretation of a given component can be problematic. Also, testing the null of non-stationarity of the first LPC leaves one vulnerable to the standard criticisms concerning the low power attached to unit root tests making it difficult to reject the null of non-stationarity. A final caveat concerns a situation where there exist two or more common trends under the null hypothesis. The ADF unit root test is conducted on the series with the largest sum of squares. However, if we take equation (6), the simple Dickey-Fuller statistic is asymptotically proportional to $\sum z_{t-1}e_t \Big/ \left(\sum z_{t-1}^2\right)^{0.5}$. It is possible that the size of the test under such a null may actually be less than 5%.

IV. Results

Before proceeding to the LPC-based tests, we may first consider the traditional test for absolute beta convergence among Latin American countries. Using the data set, the following result was obtained using OLS for the full sample of sixteen Latin American countries across the study period 1960-2000:

$$\gamma_{i,t,t+T} = \underset{(0.047)}{0.041} - \underset{(0.006)}{0.004} y_{i,t} + \varepsilon_{i,t}, \tag{7}$$

[6] If first LPC is stationary, this will imply that all real per capita incomes within the sample are stationary. Although there are no common trends among the $n+1$ countries, this result would at least imply non-divergence among the series. As is seen later, this particular state of affairs is precluded by the unit root tests.

where γ is the annualised growth rate of per capita GDP over T time periods, y_t is the initial level of per capita GDP of country i and standard errors are reported in parentheses. While the negative slope conforms to the priors suggested by the traditional approach, it is insignificantly different from zero. Thus using traditional convergence tests applied to this data set yields results that are not supportive of convergence.

Table 1. DFGLS unit root tests on per capita income

Country	Group	1960-2000		1981-2000	
		No trend	Trend	No trend	Trend
Argentina	LAIA	-1.174	-2.327	-0.945	-1.877
Bolivia	LAIA	-1.847 *	-2.286	-0.496	-0.658
Brazil	LAIA	0.916	-0.961	-0.614	-3.670 ***
Chile	LAIA	1.537	-1.198	-0.268	-1.555
Columbia	LAIA	-0.174	-1.450	-1.222	-2.293
Costa Rica	CACM	-0.557	-1.316	-0.671	-1.107
Ecuador	LAIA	-0.373	-0.339	-0.260	-2.073
El Salvador	CACM	-0.564	-2.394	-0.763	-1.328
Guatemala	CACM	0.020	-1.214	-1.460	-1.465
Honduras	CACM	-0.624	-0.987	-0.540	-3.103 **
Mexico	LAIA	1.057	-1.060	-0.690	-0.862
Nicaragua	CACM	0.140	-1.171	0.479	-1.509
Paraguay	LAIA	-1.097	-2.065	-0.954	-1.057
Peru	LAIA	-0.656	-1.287	-1.507	-2.153
Uruguay	LAIA	0.414	-1.695	-1.570	-1.904
Venezuela	LAIA	-0.415	-1.587	-0.412	-1.876

Note: These are DFGLS unit root tests advocated by Elliot, Rothenberg and Stock (1996). In all cases, the lag lengths are selected on the basis of the Akaike Information Criteria (AIC). Excluding a trend, ***, ** and * indicate rejection of the null of non-stationarity the 1, 5 and 10% levels using critical values of -2.58, -1.95 and -1.62 respectively. Including a trend, the 1, 5 and 10% critical values used are -3.48, -2.89, -2.57 respectively.

In the search for long-run relationships among the real per capita incomes using the alternative LPC-based approach, we first require that the series are non-stationary. An important issue to consider is whether or not the use of transformed data by means of natural logarithms is appropriate. For this purpose, the series were subjected to a range of tests advocated by Franses and McAleer (1998) based on tests of non-linear transformations on ADF auxiliary regressions.[7] At the 5% significance level, these tests indicated that in the case of El Salvador (1960-2000), the standard ADF regressions are not appropriately specified to test for a unit root in y. For this study period, El Salvador's per capita GDP is not transformed into natural logarithms. Table 1 reports DFGLS unit root tests advocated by Elliot, Rothenberg and Stock (1996) for all the countries. These are unit root tests that offer more power than the standard ADF tests and at the 5% significance level, the null of non-stationarity is only rejected in the cases of Brazil (1981-2000) and Honduras (1981-2000). The tests for convergence are conducted with these countries included and then excluded to judge whether the qualitative outcomes of the tests are affected.

Table 2. Stationary of the first LPC based on differences

Group	Period	n	DFGLS - no trend	DFGLS - trend	Base County
ALL	1960-2000	15	-1.202	-2.074	Ven
ALL	1981-2000	15	-1.428	-1.337	Ven
ALL (excl. Bra and Hon)	1981-2000	13	-1.319	-1.190	Ven
CACM	1960-2000	4	-2.156**	-2.521	Nic
CACM	1981-2000	4	-0.070	-2.712*	Nic
CACM (excl. Hon)	1981-2000	3	-0.059	-2.744*	Nic
LAIA	1960-2000	10	-0.612	-0.303	Ven
LAIA	1981-2000	10	-1.337	-2.281	Ven
LAIA (excl. Bra)	1981-2000	9	-0.694	-1.390	Ven
LAIA	1960-1980	10	-0.556	-1.757	Ven
North	1960-2000	8	-1.202	-2.074	Ven
South	1960-2000	6	-0.659	-0.691	Uru

Note: These are DFGLS unit root tests conducted on the first largest principal component (LPC) based on n real per capita income differences with respect to the designated base country. Details on lag length selection and critical values are given in Table 1. ** and * indicate rejection of the null of non-stationarity at the 5 and 10% levels of significance.

[7] The ADF auxiliary regression is given by $\Delta y_t = \mu_0 + \mu_1 time + \beta y_{t-1} + \gamma_1 \Delta y_{t-1} + \lambda(\Delta y_{t-1})^2 + \varepsilon_t$. The two null hypotheses of interest are $\lambda = 0$ and $\beta = 0$. Throughout, both nulls were generally accepted. However, in the case of El Salvador, both nulls were rejected at the 5% significance level suggesting that it is inappropriate to engage in a non-linear transformation of this output series.

The first stage of the convergence test is to take each of the three groupings and express per capita income with respect to a chosen base country. The choice of base countries includes Venezuela for the LAIA countries, Nicaragua for the CACM countries and Venezuela for the grouping that comprises all countries. Table 2 reports that in all cases except the CACM countries, the null of non-stationarity of the first LPC is accepted at the 10% significance level. Strong convergence with homogeneity is generally rejected for the LAIA and indeed, all Latin American countries and we are therefore unable to conclude that the movement in LDC per capita income levels are characterized as being convergent in the long-run with a coefficient of unity. This evidence of non-convergence also applies to the LAIA countries prior to 1981 (1960-80) as well as the geographical groupings based on North and South.

Table 3. Stationary of the LPCs based on real per capita income levels

Group	Period	n+1	LPC	DFGLS - no trend	DFGLS - trend	k
ALL	1960-2000	16	3	-2.021**	-2.578*	2
ALL	1981-2000	16	5	-2.984***	-2.784*	4
ALL (excl. Bra and Hon)	1981-2000	14	4	-3.213***	-3.262**	3
LAIA	1960-2000	11	5	-2.009**	-2.747*	4
LAIA	1981-2000	11	2	0.280	-3.244**	1
LAIA (excl. Bra)	1981-2000	10	2	0.280	-3.268**	1
LAIA	1960-1980	11	3	-2.025**	-2.231	2
North	1960-2000	9	3	-2.039**	-2.618**	2
South	1960-2000	7	5	-3.119***	-4.096***	4

Note: The column headed n+1 refers to the number of countries. The column headed LPC indicates which LPC is the first that is identified as being stationary according to the DFGLS unit root tests. Details on lag length selection and critical values are given in Table 1. ***, ** and * indicate rejection of the null of non-stationarity at the 1, 5 and 10% levels of significance respectively in the unit root tests. The column headed k indicates the number of common shared trends present for each group.

The second stage of the convergence test applies to those groups for whom the first LPC was non-stationary. This second test is based on the search for a single common trend among the series in levels form rather than differences with respect to base country. The results reported in Table 3 suggest that it is only in the case of the LAIA countries (1981-2000) that a single common trend is confirmed where the second LPC is the first principal component that is stationary. However, since the first stage of the test found against strong convergence, we conclude that homogeneity with respect to long-run movements in income levels is not present here and so the LAIA group is characterized as being weakly convergent. However, this evidence of weak convergence for the LAIA countries does not extend across the full study period of 1960-2000 where four single common trends are present among the eleven LAIA countries,[8] or over the sub-period 1960-1980 where two common trends are present. Evidence of multiple common shared trends is also present in the case of all the Latin American countries together as well as the geographical groupings based on North and South. Overall, the firmest evidence in favor of weak convergence in Table 3 is associated with the LAIA countries during the period 1981-2000 since the consideration of alternative groups and alternative sub-periods points towards the presence of multiple common trends. These latter findings are in principle consistent with Dobson and Ramlogan (2002) who do not find convincing evidence in favor of sigma convergence for their 1960-90 study period of Latin American countries.

The results reported in Tables 2 and 3 indicate that evidence of convergence is strongest in the case of the CACM rather than LAIA. It should be remembered that many Latin American economies have experienced serious turbulence during the study period.[9] It is therefore pertinent to ask whether the results obtained for the LAIA in Tables 2 and 3 are sensitive to structural breaks that lead one to find in favor of non-stationarity. To address this, one may employ unit root tests advocated by Perron (1997) that endogenously determine structural breakpoints in the LPCs. Using these tests, Table 4 reports that we are still unable to reject the non-stationary null at the 5% significance level in all cases. Therefore, even allowing for the abovementioned turbulence, we still find that the first LPC based on income differentials and the second LPC with respect to income levels are both non-stationary.[10]

[8] In the case of the LAIA sample (1960-2000), the fifth LPC is the first principal component that is stationary. This suggests there are seven cointegrating vectors and therefore four single common trends among the eleven LAIA countries.

[9] For example, Brazil has contended with oil price shocks, the external debt crisis and the effects of stabilisation plans. To many, the first half of the 1981-2000 sub-period is referred to as the 'lost decade'.

[10] In the latter case, stationarity of the second LPC would have implied a single common trend among the LAIA countries for the periods 1960-2000 and 1960-1980.

Table 4. Perron (1997) unit root tests on LPCs

Group	Period	IO2	AO
(a) First LPC based on real per capita income differences			
LAIA	1960-2000	-4.568	-3.358
		(1971Q1)	(1986Q1)
LAIA	1981-2000	-4.860	-4.712
		(1994Q1)	(1992Q2)
LAIA (excl. Bra)	1981-2000	-4.462	-4.645
		(1994Q1)	(1994Q1)
LAIA	1960-1980	-3.670	-3.742
		(1969Q1)	(1967Q1)
(b) Second LPC based on real per capita income levels			
LAIA	1960-2000	-4.592	-3.335
		(1971Q1)	(1986Q1)
LAIA	1960-1980	-3.671	-3.767
		(1969Q1)	(1969Q1)

Note: These are Perron (1997) unit root tests based on endogenously-determined structural breakpoints (given in parentheses). IO2 denotes tests that incorporate an innovational outlier with a change in the intercept and in the slope, and AO denotes tests that incorporate an additive outlier with a change in the slope only but where both segments of the trend function are joined at the time break. With respect to the null of non-stationarity, the 10% critical values are -5.59 and -4.83 in the cases of IO2 and AO respectively.

With respect to the results reported in Tables 2 and 3, it is interesting to test for the speeds of adjustment towards convergence. Strong convergence is identified in the case of the CACM countries so in the case of the first LPC for the full 1960-2000 period, we have

$$\Delta LPC_{1t} = const - 0.133 LPC_{1,t-1} + lags + residuals, \tag{8}$$

where LPC_1 denotes the first LPC. Using this result, the half-life of a deviation from stationarity with respect to LPC_1 is computed as $[\ln 0.5 / \ln(1-0.133)] = 4.857$ years. This is faster than the oft-cited 2% convergence rate that is quoted elsewhere in the literature in connection with beta convergence. However, it should be remembered that rather than testing for beta convergence, this paper is testing an alternative notion of convergence where, in the convergent state, per capita GDP's move together over time. This does not necessarily mean that poorer countries have caught up with richer countries. In the case of the LAIA countries (1960-2000), Table 3 reports that the fifth LPC is the first that is found to be stationary. In this case, we have

$$\Delta LPC_{5t} = const - 0.287 LPC_{5,t-1} + lags + residuals, \tag{9}$$

where LPC_5 denotes the fifth LPC where incomes are expressed in levels rather deviations from base country. Using this particular result, the half-life of a deviation from stationarity with respect to the fifth LPC is computed as $[\ln 0.5/\ln(1-0.287)] = 2.049$ years. This half-life is considerably shorter than for the CACM countries but, of course, applies to a far weaker notion of convergence because there are four single common shared trends among the eleven LAIA countries.

Using the data for income differentials defined with respect the chosen base countries, one may follow the alternative approach pursued by McCoskey (2002), in her study of convergence in sub-Saharan Africa, based on panel data unit root and cointegration testing. The IPS panel data unit root test advocated by Im et al. (2003) employs a test statistic that is based on the average ADF statistic across the sample using demeaned data for $(y_i - y_G)$. The null hypothesis specifies that *all* series or differentials in the panel are non-stationary against an alternative that *at least one* series or differential is stationary. These hypotheses are clearly different from those implied through testing the stationarity of the first LPC where the null is that at *least one* differential is non-stationary against the alternative that *all* differentials are stationary. Rejection of the null is this case offers a much stronger notion of convergence than under IPS because in the latter case, it might simply be the case that as few as one differential is responsible for rejecting the non-stationary null.

Table 5 reports that the IPS panel data unit root test rejects the null at the 5% significance level in the case of the CACM countries (1981-2000). In all other cases, the null is accepted at the 5% significance level. These results are consistent with the results reported in Table 2. However, it should be pointed out that unlike testing the stationarity of the first LPC, the panel data unit root test *is* sensitive to the choice of base country and it is possible that acceptance of the null under the IPS test may simply be due to the choice of base country. Table 5 also reports the findings from the earlier LL panel data unit root test advocated by Levin and Lin (1993). This test offers very restrictive joint and null hypotheses where *all* members

Table 5. Panel data unit root tests

Group	Period	*n*	Base County	LL	IPS
ALL	1960-2000	15	Ven	0.220	-0.263
ALL	1981-2000	15	Ven	-0.212	-0.585
ALL (excl. Bra and Hon)	1981-2000	13	Ven	-0.112	-0.614
CACM	1960-2000	4	Nic	-1.017	-1.636 *
CACM	1981-2000	4	Nic	-1.452 *	-1.757 **
CACM (excl. Hon)	1981-2000	3	Nic	-1.656 **	-1.800 **
LAIA	1960-2000	10	Ven	0.108	-0.605
LAIA	1981-2000	10	Ven	-0.518	-1.240
LAIA (excl. Bra)	1981-2000	9	Ven	-0.537	-1.416 *

Note: LL and IPS denote the Levin and Lin (1993) and Im, Pesaran and Shin (2003) panel data unit root tests. Individual lag lengths are based on the AIC. Both statistics are distributed as standard normal as both N and T grow large. ***, ** and * denote rejection of the null of joint non-stationarity at the 1, 5 and 10% significance levels with critical values of -2.33, -1.64 and -1.28 respectively.

of the panel series are either non-stationary or stationary with common autoregressive parameters. At the 5% significance level, the LL test confirms convergence with respect to the CACM countries over the period 1981-2000.

With respect to panel data cointegration testing, Table 6 reports Pedroni (1999) panel data cointegration tests for long-run relationships between per capita GDP and the designated base countries. G*roup PP* is a non-parametric statistic that is analogous to the Phillips-Perron *t* statistic and G*roup ADF* is a parametric statistic and analogous to the ADF *t* statistic. This latter statistic is analogous to the Im, Pesaran and Shin (2003) test for a unit root panel that is applied to the estimated residuals of a cointegrating regression. *Group PP* and *Group ADF* are referred to as *between-dimension* statistics that average the estimated autoregressive coefficients for each country. Both these tests are asymptotically normal and the results offer mixed evidence concerning the extent of convergence. Following Pedroni (2001), one may estimate the long-run relationship between y_i and y_G through dynamic ordinary least squares estimation. Depending on whether the stationary series are included or excluded, the slope coefficient ranges from zero to unity. Again, these results may be sensitive to the choice of base country where rejection of the non-cointegration null may be attributable to the presence of just a single cointegrating relationship from within the panel.

Table 6. Panel data cointegration tests

Group	Period	n	Base	Group PP	Group ADF	β	$t_{\beta=0}$	$t_{\beta=1}$
ALL	1960-2000	15	Ven	4.853	2.765	N/A	N/A	N/A
ALL	1981-2000	15	Ven	-2.026**	-2.900***	0.099	-0.323	-7.096***
ALL (excl. Bra and Hon)	1981-2000	13	Ven	-1.837**	-2.024**	0.481	4.540***	-0.688
CACM	1960-2000	4	Nic	0.684	-0.847	N/A	N/A	N/A
LAIA	1960-2000	10	Ven	3.869	1.807	N/A	N/A	N/A
LAIA	1981-2000	10	Ven	-0.866	-2.853***	0.548	1.401	-3.954***
LAIA (excl. Bra)	1981-2000	9	Ven	-0.842	-2.206**	0.834	3.931***	-0.500

Note: The columns headed Group PP and Group ADF are Pedroni tests for cointegration between y_i and y_G where ***, ** and * denote rejection of the null of joint non-cointegration at the 1, 5 and 10% significance levels (see Table 5 for critical values). Where the null of non-cointegration is rejected, Column 7 reports the estimated slope (β) and Columns 8 and 9 report t statistics for the null of a zero and then unity slope. Individual lag lengths are based on the AIC. N/A indicates where the null of non-cointegration is accepted. In these cases, it is inappropriate to report long-run slope estimates and associated t-statistics.

V. Conclusion

This paper has tested for economic convergence among Latin American countries- a relatively unexplored area- using groupings based on key agreements concerning trade liberalization and cooperation. For this purpose, convergence is addressed in an alternative way through the application of principal components and cointegration analysis. This multivariate technique has advantages over existing methods because less demand is placed on limited data sets and the qualitative outcome of the test is invariant to the choice of base country. In addition to this, this technique offers a different perspective on convergence based on the co-movements of real per capita outputs rather than the traditional sigma and beta convergence. There is evidence that convergence is most likely to be found within convergence clubs based on trade agreements. Using a sample of sixteen Latin American countries, we find that strong long-run convergence is only confirmed in the cases of the Central American Common Market countries over the period 1960-2000. However, a weaker form of convergence is applicable in the case of the Latin American Integration Association over the period 1981-2000. Such evidence is not present when we consider all Latin American countries considered together or groupings that are based on geographical location. The implications of our findings are twofold. First, it is not necessarily the case that convergence is restricted to smaller groups of LDCs. For example, we are able to identify the presence of a single common shared trend driving the eleven countries taken from the Latin American Integration Association. Second, groupings and sub-periods that exhibit little or no evidence of convergence provide a case for additional regional development policies aimed at facilitating closer integration among member states. Bearing in mind the findings from this study, several avenues for future research are brought to light. Researchers may reflect on why some international agreements on increased cooperation are more conducive towards convergence than others. Future research may also reflect on alternative measures of long-run convergence perhaps utilizing improved panel data techniques that enable the researcher to identify which panel members are responsible for rejecting non-stationary or non-cointegrating null hypotheses.

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