TRADE OPENNESS AND INCOME:

A TALE OF TWO REGIONS

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Abstract. In this article we present evidence of the long-run effect of trade openness on income per worker for two regions that have followed different liberalization strategies, namely Asia and Latin America. A model that re-examines these questions is estimated for two panels of Asian and Latin American countries over the 1980-2008 period using a novel empirical approach that accounts for endogeneity as well as for the time series properties of the variables involved. From an econometric point of view, we apply recent panel cointegration techniques based on factor models that account for two additional elements usually neglected in previous empirical literature: cross-dependence and structural breaks. The results point to a positive impact of trade openness in both Asia and Latin America although the size is smaller in the second region. We associate this finding with the degree to which trade was managed in both regions of the developing world.

***Keywords*:** GDP per worker, trade openness, panel cointegration, structural breaks, cross-section dependence, Asia, Latin America.

***JEL classification*:** F15, F43, C22, O40.

1. Introduction

This paper follows the tradition of panel studies looking at the link between trade openness and standards of living as proposed by Frankel and Romer (1999) and Alcalá and Ciccone (2004). Although there is an extensive empirical literature on the trade-income link[1] we think it is still worthwhile to re-examine this relationship. First of all, the issue whether trade openness promotes development in terms of income generation has not been settled yet in empirical studies, specially taking into account the possible non-stationary nature of the variables involved in this relationship in a time dimension. Second, the use of real trade openness[2] as an outcome variable (instead of trade policy measures which are extremely difficult to be observed over longer periods of time) might still allow to gain important insights as to the impact of trade on GDP per worker (Alcalá and Ciccone, 2004; Doyle and Martínez-Zarzoso, 2011).

Nonetheless, it is generally known that the Latin American (LA hereafter) and the Asian region followed quite different trade regimes over the last 30 years (Chen, 1999; Narula, 2002; Duran, Mulder and Onodera, 2008). While Latin America concentrated on fast trade liberalization policies, Asia followed a more gradual approach and pragmatic trade policies with much more government intervention. After the severe LA debt crisis that began in 1982, the LA countries were forced to follow structural adjustment programs that were composed of trade liberalization, macroeconomic stabilization, deregulation and privatization. Around 1985 basically all LA countries had moved to a strategy of (unilateral)[3] trade liberalization implying the drastic reduction or abolition of non-tariff and tariff trade barriers. Import tariffs were rapidly cut and export taxes were slashed[4]. The idea was to stop the anti-export

[1] See, among others, Edwards (1992, 1993, 1994 and 1998), Rodríguez and Rodrik (2001), Baldwin, (2003), Dollar and Kraay (2003), Lee et al. (2004) and Singh (2010).
[2] The reasons why real openness should be used instead of nominal openness are detailed in Acalá and Ciccone (2004).
[3] Existing rigidities due to a long tradition of import substitution policies (in the 1960s and 1970s) were to be abolished rapidly even without trade liberalization concessions from the trading partners.
[4] Chile started unilateral trade liberalization already in 1973 but the majority of Latin American countries followed this strategy only at the beginning of the 1980s.

bias of the previous import substitution regime and to create a neutral trade regime in which trade flows should be determined by comparative advantage (Edwards 1993, 1994). There was heated debate in academics on whether trade liberalization should be abrupt or gradual. The LA countries mostly followed the 'one shot' strategy, which forced economic agents to adjust in a speedy way. This of course caused sudden adjustment costs but also changed the incentive structure in a clear and visible way. Opponents of the strategy called this the shock therapy, but this strategy was mostly applied with the intention to avoid a building up of resistance against trade liberalization. Edwards and Edwards (1986) also pointed to the importance of the sequencing of the liberalization measures. Macroeconomic stabilization and trade liberalization had to precede the liberalization of the domestic capital market and the capital account. The opening of the capital account favors the appreciation of the real exchange rate and thus counteracts the development of export trade.[5] By keeping the real exchange rate at a realistic and competitive level, the sectors with comparative advantage could become competitive and the sometimes heavily taxed export industry be put into a position to compete on world markets. In Asia, in contrast, trade policy did not follow the pure free market approach, which was propagated in LA. Specific industries were promoted through government inducement with subsidized loans and tax incentives for investment. In particular, a low-tech industry was set up covering textiles and clothing, footwear and toys to be followed by a medium and high-tech industry covering non-electrical and electrical machinery, chemicals, pharmaceuticals and information technology. The creation of a high-tech industry was especially successful in Japan and South Korea. The government played an important guiding and coordinating role in strengthening and building up a highly competitive export industry. Governments in Asia set up export processing zones and generously granted export subsidies and export credits (Hwang Doo-yun, 2001; Rodrik, 2006). Equally important were the governmental incentives designed to remove the obstacles

[5] Argentina did not follow this sequencing strategy. It liberalized the trade account and the capital account simultaneously in 1976 and had to give up its first liberalization attempt in 1978.

to private investment and to improve the investment climate (Rodrik, 2011). It is important to note that while in Asia industries were supported only as long as they exported, this approach was lacking in LA, where the "infant" industry argument were disregarded and liberal trade policies were propagated without an accompanying industrial policy and without the necessary state support to energize the private sector.

From an economic history perspective, we could imagine a structural break in the export series in the mid-1980s in LA (due to the structural adjustment programs) and in the late 1990s in Asia (due to the burst of a housing bubble and a current account crisis in 1997 which triggered harsh economic reforms).

Acting under different trade regimes, LA and Asia had quite similar exchange rate regimes. In most LA countries pegged exchange rates were given up at the end of the 1990s or the early 2000s, starting with Brazil in 1999. In South-East Asia the fixed exchange rate systems were abandoned after the debt crisis in 1997 and floating exchange rate regimes followed. The degree of floating varied from country to country. E.g. China applied a managed float, trying to keep an undervalued exchange rate. However, the real exchange rates that accompanied the trade liberalization process were more competitive and stable in Asia and overvalued in LA, according to Duran et al. (2008).

Summarizing, although both regions had pursued a trade liberalization strategy with the same main aim, the approach has been substantially different. As a consequence, Asia has been more successful in diversifying its export structure, whereas LA has remained focused on commodities and low value-added products.

The objective of the study is twofold: First, to uncover the relationship between real trade openness and GDP per worker (the so-called labor productivity) utilizing recent panel cointegration techniques and second, comparing the LA to the Asian experience. Current panel cointegration techniques allow us to account for possible structural breaks and cross-section dependence of the series. Furthermore, the use of these techniques is handy for several reasons: it enables us to control not only for endogeneity but also for the autocorrelation of the error terms and the omitted variables problem. To the best of our knowledge, estimators robust to cross section dependencies and structural breaks have never been applied before to the estimation of the income-openness link. As for the comparison of LA and Asia, this exercise might show whether increases in income are rather a function of the free interplay of markets or government intervention in the form of industrial policy/ strategic trade policy.

The main results show that after controlling for endogeneity and structural breaks, openness has a positive and significant effect on the level of income in both regions LA and Asia and the effect is higher in magnitude for the second region. The effect is magnified for the Asian sample when excluding Japan. We find structural breaks for Asia in 1997 and for LA 1985.

The paper is organized as follows. Section 2 discusses the related empirical literature on growth and trade, presenting an econometric approach based on panel cointegration that overcomes some of the problems usually present in the current literature. Section 3 describes the data and discusses the empirical results. A final section concludes.

2. Theoretical and Empirical approaches to the trade-income link: Where do we stand today?

The theoretical channels through which trade openness impacts on per capita income are basically three: the neoclassical, the endogenous growth and the institutional approach.

First, in the neoclassical approach, trade patterns among countries are determined by its comparative advantage, according to which each country exports goods produced with a lower relative unitary-costs than its competitors and thus maximizes its welfare by concentrating on those activities in which it is most economically efficient[6]. Generally, according to the neoclassical theory, an opening up to trade does not lead to a long-run increase in the rate of growth, only to an increase in the level of income. Second, the endogenous growth approach found that trade openness could impact on both the level of income and the long-run rate of growth of an economy through scale, allocation, spillover and redundancy effects[7]. Finally, although the impact of institutions has not yet been incorporated into economic growth theory, it is widely recognized that without basic institutions (law and order, well-defined property rights and impartially-enforced contracts), the expected positive response to trade openness may not appear.

In the past decades empirical research on the link between openness and growth was extraordinarily abundant. Nonetheless the research on the question of whether more open economies grow faster or produce a higher per capita income has not been settled yet.

Michaely (1977), Krueger (1978) and Bhagwati (1978) started off the debate on whether foreign trade regimes matter for growth. These authors found empirical evidence supporting that outward oriented countries outperformed inward oriented countries in terms of growth. This finding was based on theoretical reflections and received an empirical underpinning by "looking" at factual criteria of trade liberalization, which led to a classification into outward and inward-oriented regimes on the one hand and into good and poor growth performers on the other. Later, Balassa (1985) was among the first economists who used econometric analysis to study the relationship between openness and growth. Cross-country regression analysis revealed a positive link between trade openness and growth. Reliance on cross-

[6] The gains from trade may be static (such as improvements in allocation efficiency of resources) or dynamic through imported technology or learning-by-doing effects.

[7] Scale effects are derived from the closer integration of a country into the world market, while allocation effects arise from the accumulation of factors of production such as human or physical capital or R&D and then benefit those sectors that use these factors intensively. Poor access to imported capital goods embodying improved technology is considered to be a particularly growth-inhibiting factor. The spillover effect is a related effect with trade leading to the diffusion of new knowledge. Similarly, open trade leads to a reduction of unnecessary duplication of research, eliminating redundancy in R&D.

country studies was intensified in the 1990s.[8] Researchers looked at a whole bundle of countries (developing and developed) and their average growth rates in a certain period (e.g. from 1960-1999) and the role played by openness. Openness was measured by different indicators, such as average tariff levels, import value of tariffs, existence of state-run monopolies, black market premium, share of trade in GDP etc. This led to vehement discussions of whether it is more relevant to look at trade policy measures (such as change in tariffs, share of tariff covered imports in total imports) or at criteria that are related to the efficiency of markets, such as interference and power of the state (state-run enterprises), existence of macroeconomic distortions, or black market premium[9] (see Dollar, 1992; Ben-David, 1993; Sachs and Warner, 1995; Edwards, 1998; Frankel and Romer, 1999). The prominent evidence of these studies pointed to a positive and significant relationship between a more liberal trade policy/a higher trade volume and economic growth. However, this evidence was challenged by either a critique on the pitfalls and arbitrariness of the trade measures used or by questioning the econometric techniques used (Rodríguez and Rodrik, 2001). Concerning the econometric techniques, the main flaw associated to cross-country analyses was their inability to deal with unobserved heterogeneity, i.e. country characteristics, such as institutions, governance, mentalities and so forth. Second, most of the studies paid not enough attention to the issue of endogeneity[10]. As a result of the above criticisms, in a second wave of empirical studies it became state of the art to estimate the openness-growth nexus working with panel data (Harrison, 1996). Most studies have used fixed effects (within estimation) to capture time-invariant country characteristics and to control for the time-invariant part of unobserved heterogeneity (Wang et al., 2004; Felbermayr, G.J., 2005; Doyle and Martínez-Zarzoso, 2011). Endogeneity was taken care of by instrumenting for trade openness through geographic and historical variables (distance between trading partners, distance to the equator, colonial history, and settler mortality).

[8] Barro (1991) established the cross-section growth literature that did not include trade openness at that time.
[9] Which reflects the difference between the official exchange rate and the exchange rate in the black market.
[10] Hsiao (1987) wrote an influential paper on the causality and exogeneity between exports and economic growth.

Two types of GMM estimation[11] became popular in later years when the regression equation models were set up as dynamic panel data models. First, difference GMM (Arellano and Bond, 1991), in which the equation is estimated with variables in first differences to eliminate the fixed effects, and the endogeneity of the lagged dependent variable is tackled by instrumenting the endogenous variables using its lagged first differences (or levels). Second, system GMM, which consists in estimating a system of two equations, the first equation in first differences (using the lagged variables of the endogenous variables in levels as instruments) *and* the second in levels (using the lagged first differences of the endogenous variables as instruments), was the recommended option if the series were persistent. This was the approach proposed by Blundell and Bond (2000). However, neither the difference GMM nor the system GMM estimators would be right approaches if the dependent variable contains a unit root. Nowadays this approach is criticized for using weak instruments, i.e. the correlation between original variable in levels and lagged first differences of variables is considered too low. For the Asian case, using system GMM approach, Das and Paul (2011) obtained a positive and significant link between openness and growth. Rodriguez (2007) claims that when one introduces several measures of geography in the regression, the coefficient on trade becomes statistically insignificant, while others, like Harrison (1996), Winters (2004) or Doyle and Martínez-Zarzoso (2011), reach more optimistic conclusions. In particular, the latter authors' results suggest a robust positive relationship between real openness and labor productivity from the 1990s onwards[12].

There are also a number of papers that focus on specific regions. In Nelson and Zolnik (2013) the effect of openness on income is estimated using panel-data and instrumental variable techniques for several regions, among them Latin America and the Caribbean, for which they report a positive and significant coefficient for the period 1980-2005. For the Asian region,

[11] Which goes back to Hansen (1982).
[12] System-GMM dynamic panel estimation methods are used to address potential biases associated with cross-section estimations such as omitted variable problems and endogeneity of explanatory variables.

the paper by Das and Paul (2011) has been already mentioned before and in Mustafa, Rizov and Kernohan (2013) a similar results was obtained using also GMM techniques (openness had a positive and significant effect on income with an estimated coefficient of around 0.18).

A more descriptive paper, which focuses on the two-regions case, namely LA and East Asia (EA) is Duran et al (2008). The authors highlight the differences between the trade policy approaches –fast track in LA and gradual in EA- followed in both regions as one of the main factors explaining their differential development paths. They concluded that the more successful growth experience in EA in terms of exports and income in comparison with LA is explained by a number of factors, namely the sequencing and content of trade reforms, the level of macroeconomic and exchange rate stability and the level of integration in the global economy. However, the authors do not embark on any econometric exercise and limit their research to a comparison between both development experiences. Notwithstanding, the evidence presented will be valuable to interpret the results in the present research. A paper that also focuses on the EA-LA case is Chen (1999). The author specifically tests the relationship between trade openness and economic growth using a growth model based on Romer (1990), which is applied to long-run averaged data over the period 1970-1992 for a sample of 34 countries in the area. The main results indicate a positive and significant effect of trade, measured with five proxies, on income growth. One of the proxies is a size-free trade-GDP ratio. However, the authors fail to estimate different slope coefficients for LA and EA countries and use only a dummy for the EA countries to differentiate both regions. Moreover, the author fails to control for the time series properties of the data. We overcome these shortcomings in this paper by allowing for regional-specific coefficients for all the determinants in our model and using a more comprehensive empirical approach, which exploits the time series properties of the data and allows for structural breaks and cross-section dependence. Indeed, a more recent criticism applied to the studies on the trade-income link came from time series economists. They argued that not accounting for the underlying time series properties might result in the estimation of spurious relationships and that, after all, the focus of up-to-date studies should be on the long run (rather than the short- to medium run). A further criticism was that the omitted variable problem had been neglected in the standard (static and dynamic) panel data regressions. The only steps taken to alleviate this problem so far have been an increased usage of fixed or random effects which, however, capture only time-invariant omitted variables. However, controlling for time-invariant omitted variables is not always enough to avoid the above-mentioned problem and a proper treatment of the time dimension is desirable. While the heterogeneity bias is controlled through the use of fixed-effects, a second kind of misspecification is related to dynamics.

This criticism led to a new range of empirical studies in which the time series properties of the variables started to play a very important role and in which emphasis is put on whether the series are co-integrated. According to Herzer (2013) a reduced form model is justified provided cointegration between the variables holds. The estimated trade-income coefficient can then be estimated with enhanced estimation techniques and will be unbiased and consistent even if additional variables[13] are added to the bivariate regression. Studies in this line of research are Singh (2011), Herzer (2011), Bajwa and Siddiqi (2011) and Dreger and Herzer (2013). While Singh (2011) uses time series cointegration techniques for the case of a single country (Australia), Herzer (2011), Bajwa and Siddiqi (2011) and Dreger and Herzer (2013) use panel cointegration techniques for several countries. Although Bajwa and Siddiqi (2011) is an interesting study as it provides positive support to the openness-productivity nexus for the Asian economies, it fails to control for cross-section dependence. In Herzer (2011) stationarity tests are based on individual unit-root tests (Pesaran, 2007) and cointegration tests are based on individual vector error correction models (Breitung, 2005). More recently, Dreger and Herzer (2013) focus on the study of the export-led-growth

[13] Such as investment, human capital, natural resource endowment, institutions, trade costs, distance from the equator and so forth.

hypothesis refining the econometric technique through the implementation of the error correction model (ECM) cointegration test suggested by Gengenbach et al. (2008). Even if these studies account for the existence of cross-section dependence and heterogeneity, they fail to consider the existence of structural breaks in the relationship. Recently, Falvey et al. (2012) and Henry et al. (2012) have stressed the importance to account for the existence of nonlinearities or discontinuities in the openness-productivity relationship. In these studies, they apply threshold techniques; however, a relevant issue in this analysis is that nonlinearity and instability generally are difficult to distinguish and both are compatible. Particularly, the instability in a relationship could lead to nonlinearity, and vice versa. In fact, the existence of discontinuities can be assimilated to a local approximation of a nonlinear relationship. The lack of control for structural breaks in the series may be reflected in the parameters of the estimated models that, when used for inference or forecasting, can lead to misleading results. This problem is especially true in the case of time series that cover different historical stages that can be subject to discontinuities. In general, structural breaks are a problem for the analysis of economic series, since they are usually affected by either exogenous shocks or changes in policy regimes. As a consequence, the assumption of stability in the long run relationship between openness and GDP would seem too restrictive. Therefore, not allowing for structural breaks would be an important potential shortcoming of the past research using cointegration techniques.

In this paper, we add to the above-mentioned literature in two different respects. First, we improve the specification adding to the basic bivariate model an extra variable with a time dimension: population[14]. This variable accounts for the evolution of the size of the country over time. The rest of usual suspects generally included in the growth-openness model are

[14] An alternative variable to measure country size is population density (population/area). However, since geographical area is time-invariant, the use of this variable will not alter the results in a non-stationary panel framework in which time variation is the main focus.

time invariant variables[15], which are assumed to be included in the deterministic components of the equation. Second, we apply an econometric methodology comprising of a range of techniques to test and estimate efficiently in a non-stationary panel framework, solving endogeneity problems as well as possible biases posed by structural breaks and cross-section dependence.

As we work with panel data (27 countries: 12 Latin American and 14 Asian countries) the issue of cross-section dependence becomes also relevant. Overlooking this potential problem could lead to two flaws in our empirical findings. First, the cointegration results could be spurious; second, the influence of trade openness on productivity or income could be overestimated. Since the Pesaran (2004) CD statistic reveals the existence of these dependencies, we claim that robust estimators should be employed. Although there are several alternative proposals formulated in the literature to overcome the cross-section dependence problem, when the dependence is pervasive –as in economic integrated areas- a recently proposed alternative is the use of factor models. This consists of assuming that the process is driven by a group of common factors, so that it is possible to distinguish between the idiosyncratic component and the common component.

Bearing the above points in mind, our empirical strategy proceeds as follows: First, we determine the order of integration of the variables through panel unit root tests; second, we test for cointegration among the integrated variables using panel cointegration tests; finally, we use the panel cointegration estimators to provide reliable point estimates.

Several panel unit roots tests have been formulated based on factor models[16]. In particular, Bai and Ng (2004) account for the non-stationarity of the series coming either from the common factors, the idiosyncratic component or from both. As regards to cointegration tests,

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[15] It is worth mentioning that institutional quality has also been widely included in the model. Although it is indeed a time-varying variable, the variation over time of this variable is very limited for many countries, in particular for those in the two regions considered in this paper. Hence, we assume that it effect is also included in the deterministic component of the equation.

[16] Namely, Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) and Bai and Ng (2004).

it is important to emphasize that a failure to account for the existence of changes in the cointegration relationship and/or the deterministic components affects inference on cointegration analysis, thus leading to wrong conclusions. The standard tests may not reject the null hypothesis of no cointegration when it is false, reducing the power of the test. Therefore, in this paper we propose the use of the tests developed in Banerjee and Carrion-i-Silvestre (2013). They generalize the approach in Pedroni (1999, 2004) to account for one structural break that may affect the long-run relationship in a number of different ways (cointegrating vector and/or deterministic components). Additionally, they address the cross-section dependence issue by using the factor model approach due to Bai and Ng (2004) to generalize the degree of permissible cross-section dependence allowing for idiosyncratic responses to multiple common factors. It is worth noting that the countries analyzed in this paper belong to two rather homogeneous samples (Asia and Latin America). This fact makes the assumption of a homogeneous break for each regional bloc rather suitable.

Finally, we employ Bai's et al. (2009) Continuously Updated (CUP) estimator, which is consistent in the presence of cross section dependence. This methodology not only efficiently estimates the coefficients but it is also based on the common factors decomposition, what assures a homogenous econometric approach.

To the best of our knowledge this is the first work that applies panel cointegration techniques with structural breaks to study the role of real openness on GDP per worker. Moreover, we estimate two separate panels, one for Asia and another for Latin America and compare the long-run behaviour of the two regions for the period 1980-2008.

3. Empirical strategy

3.1. Data description and model specification

Our basic setup is a stylized version of the Alcalá and Ciccone (2004) model, where instead of including geographical or institutional variables, we control for them through the deterministic components of our specification. It is worth to note that the above-described developments in econometrics make the use of large econometric specifications more trivial as long as the main variables under study, namely real trade openness and GDP per worker, are co-integrated.

We use annual data covering the period of 1980-2008 for LA and Asian countries. The data for income, nominal imports and exports, GDP in PPP US$, and population have been obtained from the Penn World Tables 7.0. (Heston et al., 2011)[17].

The equation of main interest relates real GDP per worker to real openness and country size (measured by population) as follows:

$$lprod_{it} = \alpha_i + \beta_1 lopen_{it} + \beta_2 lpop_{it} + \delta_i t + u_{it} \tag{1}$$

where $lprod_{it}$ is the log of income per worker in country i[18] over time periods $t= 1, 2,\ldots,T$ and countries $i=1, 2,\ldots,N$; $lopen_{it}$ denotes a measure of real openness[19] and $lpop_{it}$ denotes population in country i which represents the domestic scale of production measured as population. The $\boldsymbol{\alpha_i}$ and $\boldsymbol{\delta_i t}$ are, respectively, country-specific fixed effects and country-specific deterministic time trends, capturing any country-specific omitted factors that are relatively stable or evolve smoothly over time. Accordingly, in contrast to the cross-section and classical panel data studies reviewed above, we do not need to control for omitted variable bias by including direct proxies for geography, and institutions, since it can be

[17] The sample spans from 1980 to 2008 and includes two groups of countries: Asia and Latin America. The Asian sub-sample consists of 14 countries, namely Bangladesh (BGD), China (CHN), India (IDN), Indonesia (IND), Japan (JPN), Cambodia (KHM), Korea (KOR), Sri Lanka (LKA), Malaysia (MSY), Pakistan (PAK), Philippines (PHL), Singapore (SGP), Thailand (THA) and Vietnam (VNM). The Latin American group includes 12 countries: Argentina (ARG), Bolivia (BOL), Brazil (BRA), Chile (CHL), Colombia (COL), Costa Rica (CRI), Equator (ECU), Mexico (MEX), Paraguay (PRY), Peru (PER), Uruguay (URY) and Venezuela (VEN).

[18] Labor Productivity is defined as $\left(\frac{PPP\,GDP_{it}}{Work_{it}}\right)$

[19] Real openness is national imports plus exports (in US $) divided by national GDP in PPP US$, that is total trade relative to PPP GDP ($Trade_{ijt}/PPPGDP_i$) in logs as suggested by Alcalá and Ciccone (2004).

assumed that all these factors are absorbed into the fixed effects and/or country-specific trend terms. Finally, u_{it} denotes the error term.

Rodriguez and Rodrik (2001) criticized previous empirical evidence on two respects: first, they claimed that the indicators of openness were inappropriate and second, that the econometric techniques commonly used were questionable. However, using different measures for openness is not absent of criticisms for several reasons, and moreover, the results using different measures of openness or a combination of them (indices) does not appear to lead to more conclusive results (Rodriguez, 2007). Therefore, our approach uses a relatively simple openness variable and elaborates on the econometric techniques to solve the typical problems of endogeneity, omitted variables and simultaneity commonly present in this literature.

In our study we use the so-called "real openness" variable suggested by Alcalá and Ciccone (2004). Conventional openness measures yield downwardly biased estimates because trade-induced productivity improvements concentrate on tradables changing relative prices against non-tradables and therefore, generating a decline in the trade/nominal GDP ratio. Therefore, this variable needs to be refined using GDP in purchasing power parity (PPP). Figures 1 and 2 show the evolution over time of real openness for our sample of countries in Asia and LA. The Figures indicate that countries in Asia are more heterogeneous and with two clear outliers, namely Singapore and Malaysia, whereas the average level of openness is more homogeneous in LA, but also with two countries -Chile and Costa Rica- showing higher than average ratios. Moreover, a steady increase in openness is observed in LA starting in the mid-1980s, whereas in Asia the ratio mainly increased from the 1990s onwards.

Moreover, GDP per worker is measured by GDP in PPP divided by total labor force. The relationship between real openness and productivity is associated with the Solow model. Since this relationship could be different for economies with a differing sectoral structure, we show in Table A.1 the share of the main economic sectors in total value added for each country in our sample. The reported figures indicate that the average share of manufacturing in GDP is 17 percent in LA and 22 percent in Asia, with a few countries in the second region with higher than average shares, namely China, Indonesia, South Korea, Malaysia, Singapore and Thailand. However the shares of industry, which also includes mining, in value added are not that different.

Finally, according to Frankel and Romer (1999), country size is included in the regression model for two reasons. First, it serves as a crude proxy for the amount of trade within a country. Accordingly, the estimate of β_2 can be used to assess whether countries also benefit from within-country trade. Second, because larger countries tend to have more opportunities for trade within their borders, and therefore lower trade shares, it is necessary to control for country size in estimating the impact of international trade on income. Otherwise, $lpop_{it}$ would enter the error term, thereby inducing a negative correlation between u_{it} and $lopen_{it}$ and thus a downward bias in the estimate of β_1.

3.2. Panel unit root tests and cross-section dependence

Since the countries in the sample belong to areas with important economic relations, it is highly probable that the series show cross section dependence. Therefore, prior to other subsequent analysis we implement the test statistic proposed by Pesaran (2004) to assess whether the time series in the panel are cross-section independent. Under the null hypothesis of cross section independence the CD statistic of Pesaran (2004) converges to the standard normal distribution. This test is valid for N and T tending to ∞ in any order and that is particularly useful for panels with small T and large N. In addition, this test is also robust to possible structural breaks, which makes it especially suitable for our study. The results in the first column of Table 1 show that the null hypothesis of independence is strongly rejected and hence cross-section dependence has to be considered when computing the panel data statistics.

Once we have found evidence of dependence, we study the order of integration of the regression variables. To this aim we employ the Bai and Ng (2004) test. Their method controls for cross-section dependence given by cross-cointegration relationships. The Bai and Ng (2004) approach decomposes the variable into a deterministic part, a vector of common factors and an idiosyncratic disturbance term. For the estimated idiosyncratic component the authors propose an ADF test for individual unit roots and a Fisher-type test for the pooled unit root hypothesis ($P_{\hat{e}}$), which has a standard normal distribution.

The estimation of the number of common factors is obtained using the panel Bayes information criterium (BIC) as suggested by Bai and Ng (2002), with a maximum of six common factors. Bai and Ng (2004) propose several tests to select the number of independent stochastic trends, k_1 in the estimated common factors, $\hat{F}_t$. If a single common factor is estimated, they recommend an ADF test whereas if several common factors are obtained, they propose an iterative procedure to select k_1: two modified Q statistics (MQ_c and MQ_f), that use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution. They test the hypothesis of $k_1 = m$ against the alternative $k_1 < m$ for m starting from $\hat{k}$. The procedure ends if at any step $k_1 = m$ cannot be rejected.

The upper and lower part of Table 2 shows the results of this test for Asia and LA for the three variables considered in the analysis (GDP per worker, real openness and population). The idiosyncratic component is found to be non-stationary for labor productivity and openness, both in the case of the Asian and the LA countries. For population, in contrast, the idiosyncratic component is found to be stationary. The results of the factor component analysis point also in the direction of non-stationarity; the null hypothesis of independent stochastic trends cannot be rejected in any of the cases. Moreover, all the variables have a total of six common factors. Hence, we have enough evidence to conclude that the variables are non-stationary and that cross-section dependence is present in data.

3.3. Evidence of cointegration and structural breaks in the openness-productivity nexus

The next step in our empirical strategy is to test whether $lprod_{it}$, $lopen_{it}$ $lpop_{it}$ (productivity, openness and size, respectively) are cointegrated. To this aim, we employ the Banerjee and Carrión-i-Silvestre (2013) test. They propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector that also accounts for the presence of cross-section dependence using factor models. It is noteworthy that inference concerning the presence of cointegration can be affected by misspecification if the existence of breaks is ignored.

Banerjee and Carrion-i-Silvestre (2013) propose eight different model specifications that vary according to the way in which the deterministic components and the cointegration vector are specified.

The common factors are estimated following the method proposed by Bai and Ng (2004). They first compute the first difference of the model; then, they take the orthogonal projections and estimate the common factors and the factor loadings using principal components. In each of these specifications, Banerjee and Carrion-i-Silvestre (2013) recover the idiosyncratic disturbance terms ($\widetilde{e}_{i,t}$) through accumulation of the estimated residuals and propose testing for the null of no cointegration against the alternative of cointegration with break using the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo *t*-ratio $t^{j}_{\widetilde{e}_i}(\lambda_i)$, $j = c, \tau, \gamma$. In the model without a time trend (Model 4) j=c. In those with a stable linear time trend (Models 1 and 5) j=τ, and finally, j=γ in the models with a changing trend (Models 2, 3 and 6).

When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms.

In Table 3 we present the results of the tests for non-cointegration Z_j^* for the model with homogeneous structural breaks for the eight potential specifications discussed above. In the left-hand side, the results of the Asian countries are shown, whereas the right hand side provides the results for LA countries. Using the BIC [20] we choose Model 3 (in Table 3) in the case of the Asian countries, that is, the one that contains a constant and a trend, which is affected by a structural break. For LA, we select both, Model 6 (Table 3), including a structural break that affects not only the deterministic components but also the cointegrating vector, and again, Model 3 (Table 3). In order to test for non-cointegration, we apply the statistics based on the accumulated idiosyncratic components, Z_j^*, that follows a normal distribution. We present the tests for all possible model specifications; in all cases the null hypothesis of non-cointegration is rejected. The break is found to happen in 1996 for the Asian Countries and in 1985 (according to Model 3) or 1997 (according to model 6) for LA countries. Although the assumption of a common break for all the countries in each group might seem a little restrictive, however the geographic homogeneity of the samples is enough to find a representative common break.

Finally, given that the existence of cointegration relationships is unambiguous, we move to the next step, which is to estimate the long-run relationship. For this purpose, in the next section we will employ consistent techniques proposed by Bai et al (2009).

3.4 Estimation of the equation for the openness-productivity nexus and main results

In the previous sections we have found evidence of the non-stationarity of the variables (income and openness) and the existence of cointegration among them. The next step in the analysis is to estimate the long-run relationship among the variables. Traditional estimation methods as Ordinary Least Squares (OLS) or the Least Squares Dummy Variables (LSDV) approach present biases and inconsistencies in our non-stationary panel setting. To avoid them, the Fully Modified (FM) estimator of Phillips and Hansen (1990) and the Dynamic Ordinary Least Squares (DOLS) estimator proposed by Saikkonen (1991) and Stock and

[20] According to the BIC six factors are found in the panel.

Watson (1993) are some of the alternatives employed in the literature. These two latter estimators consistently estimate the long-run parameters and correct for autocorrelation and endogeneity under the assumption of cross section independence. However, since the Pesaran CD has revealed the existence of dependencies among the series, alternative estimators should be employed. Bai et al. (2009) have proposed a Continuously Updated estimator (CUP) to overcome the dependence problem. They consider the following model:

$$y_{it} = x_{it}^{'}\beta + e_{it} \qquad (6)$$

where for $i = 1,...,n$, $t=1,...,T$, y_{it} is a scalar,

$$x_{it} = x_{it-1} + v_{it} \qquad (7)$$

x_{it} is a set of k non-stationary regressors, β is a $k \times 1$ vector of the common slope parameters, and e_{it} is the regression error. They assume that e_{it} is stationary and *iid* across i. The pooled least squares estimator of β is as follows:

$$\hat{\beta}_{LS} = \left(\sum_{i=1}^{n}\sum_{t=1}^{T} x_{it}x_{it}^{'} \right)^{-1} \sum_{i=1}^{n}\sum_{t=1}^{T} x_{it}y_{it} \qquad (8)$$

Although this estimator is, in general, T consistent, there is an asymptotic bias due to the long-run correlation between e_{it} and v_{it}. This bias can be estimated and a panel fully-modified estimator can be developed as in Phillips and Hansen (1990) to achieve $\sqrt{nT}$ consistency and asymptotic normality. In addition, they model cross-section dependence by imposing a factor structure on e_{it}:

$$e_{it} = \lambda_{i}^{'}F_{t} + u_{it} \qquad (9)$$

where F_{it} is an $r \times 1$ vector of latent common factors, λ_i is an $r \times 1$ vector of factor loadings and u_{it} is the idiosyncratic error. If both F_t and u_{it} are stationary, then e_{it} is also stationary. In this case, a consistent estimator of the regression coefficients can still be obtained even when the cross-section dependence is ignored. In the majority of the cases, the factors F_t are unobserved. Then F_t should be estimated along with β by minimizing the objective function;

the unobserved quantities can be replaced by initial estimates and iterate until convergence through the continuously-updated estimator (CUP) for (β, F), defined as,

$$(\hat{\beta}_{Cup}, \hat{F}_{Cup}) = \arg\min_{\beta, F} S_{nT}(\beta, F) \qquad (10)$$

The estimator $\hat{\beta}_{Cup}$ is consistent for β, although it still has a bias derived from having to estimate F_t. The authors correct this bias using two fully-modified estimators. The first one directly corrects the bias of $\hat{\beta}_{Cup}$ and is denoted $\hat{\beta}_{CupBC}$. The second one makes the correction in each iteration and is denoted $\hat{\beta}_{CupFM}$.

We proceed as follows: the first step consists of filtering the variables from the deterministic components (both in Models 3 and 6 there is a structural change not only in the constant but also in the trend). Then, with these filtered variables, we use the CUP estimators to obtain the long-run parameters. In the case of Model 6 we have to estimate the coefficients before and after the break. The value of the parameters in the second half of the sample will be the sum of the two coefficients (before and after the break).

The number of common factors for the estimation is selected according to the Principal Components Factor Analysis (PCA henceforth)[21]. Starting with the group of the Asian countries, we have concluded that the best model specification corresponds to the one with one common structural break affecting both the intercept and the slope (first part of Table 4) with a break found in 1996. The estimated parameters are positive for $lopen_{it}$ and negative for the size of the country, $lpop_{it}$. We should emphasize that the bias-corrected estimators (CUP_FM and CUP_BC) show almost identical results, which is reassuring.

Concerning the LA countries, for comparative purposes we have estimated the same specification as for Asian countries, with breaks in the intercept and the slope (Model 3)

[21] This analysis does not permit to identify the common factors, although we can reasonably think of real oil prices shocks or commodity shock prices as examples of them.

together with Model 6. In Model 3, where the break is placed around 1985, we obtain significant estimates for the two variables with the CUP estimators (see Table 5, columns 4 and 5 in the first panel). The coefficient of the variable representing the size of the country is negative, which is consistent with the literature. The second panel in Table 5 shows the results of Model 6. In this case the structural break is found in 1997, a very close date to the one obtained for Asia that was 1996. However, the parameter in the second part of the sample corresponding to $lpop_{it}$ is not significant. Therefore, we discard model 6 for Latin American countries and focus on Model 3.

5. Discussion and interpretation of results

Having discussed the properties of the above-mentioned estimators, we can neglect the results produced by the Least Squares Dummy Variables and the Bai Fully Modified estimators and can just focus on the results obtained by the CUP_FM and the CUP_BC estimators (i.e. last two columns of Tables 4 and 5). Moreover, in our case, the results do not differ much.

Concerning the deterministic components of the estimated model, we find a structural break in 1985 for LA and in 1999 for Asia[22]. Furthermore, we observe that real openness and income per worker form a long-run relationship i.e. determine each other in a consistent and systematic way over the 1980-2008 period. A one percent increase in openness leads to a 0.15 percent increase in income (productivity) in Asia. This impact is not very big, but significant. In LA, again, openness does significantly influence productivity, but the value is clearly smaller (0.08). We argue that this outcome is most probably linked to differences in the trade regime and to the differential processes of trade liberalization followed in both regions (Duran et al. 2008). On the one hand, LA countries had a relatively neutral trade regime, whereas in Asia the government played an active role in shaping competitive advantage.

[22] We also estimated the model for Asian countries excluding Japan, as suggested by one referee. The structural break for Asia without Japan was found in 2001.

Asian industries (both import substitution and export industries) were promoted by deliberate government intervention. On the other hand, the trade liberalization process started earlier and was faster in LA than in Asia, and preceded the Uruguay Round, whereas the process was more gradual in Asia and took place later, mainly after the financial crisis of 1997. This means that in Asia countries started first to support a number of sectors with industrial policies and in turn liberalized trade, whereas in LA trade was directly liberalized without using first an "infant industry" strategy. In comparison to previous research, the effect of openness on income levels was estimated for non-OECD countries in Doyle and Martínez-Zarzoso (2011) using several panel-data techniques. Although the sample of countries is not comparable, nor is the period (1980-2000), it is worth to notice that the authors find a coefficient of 0.101 for non-OECD countries (0.104 for OECD) when using a system-GMM estimation technique. The magnitude of this estimate is consistent with our results, since a broader sample –of 98 countries– was used, for which the data were available. In Nelson and Zolnik (2013) the effect of openness on income was also estimated using panel-data and instrumental variable techniques for several regions, among them Latin America and the Caribbean, for which they report a positive and significant coefficient for the period 1980-2005. Although the magnitude of the effect is higher, they report separate results for intra- and inter-regional trade and the sample of countries, period considered and estimation techniques differ from ours. They report instead non-significant effects of trade on income for Sub_Saharan Africa and even negative for Europe.

As for the effect of population on productivity, it is negative both in Asia and in LA. However, the parameter is twice as big in the latter case (-0.74 against -1.38). This means that a one percent increase in population in Asia decreases labor productivity by 0.74%, whereas in LA this reduction is almost twice as large (1.38%). This result is consistent with the theory. In comparison with previous literature, Doyle and Martínez-Zarzoso (2011) found a smaller population coefficient (-0.526) using a panel fixed-effects model estimated for a sample of 85 countries over the period 1980-2000. However, the population coefficient turned out to be positive and significant when using Hausman-Taylor or system-GMM estimation technique and controlling for the endogeneity of the openness variable. These authors also included in the latter specifications (Hausman-Taylor, system-GMM) area of the country as a measure for country size, but the estimated coefficient was in most cases not statistically significant. In theory, a larger area can have a negative effect on productivity via lower intra-country trade and/or a positive impact through increased economies of scale. Focusing on country size and holding population density constant (population/area) the effect of country size on productivity would be the sum of both the log of population and the log of area coefficients (Frankel and Romer, 1999). Since we identify a negative scale effect in our results, and area is a time-invariant factor that is subsumed in the unobserved heterogeneity that is eliminated from the model, we conclude that the size or scale of a country has a net negative effect on productivity for these two regions. This is in accordance to Frankel and Romer (1999) expectation that country size impacts not only on the propensity to trade externally, but also internally. Larger countries would be expected to offer more opportunities for within-country trade and therefore tend to be less open. Nevertheless, is it worth to notice that the population coefficient turns out to be positive and significant when the model is re-estimated for the Asian sample excluding Japan (See Appendix, Tables A.2 and A.3).

6. Concluding remarks

The progress of globalization in the last three decades has drawn the attention of researchers and policy-makers, particularly in developing countries, on the relationship between economic performance and trade openness. Despite voluminous work in this area, the findings are far from being conclusive. Asian and LA countries constitute two natural case studies for a comparative exercise as they have followed different development strategies.

Using a panel cointegration approach that accounts for the existence of cross-section dependence and breaks we overcome the usual econometric flaws present in previous empirical studies using cross-section or panel data. A modified version of the model proposed by Alcalá and Ciccone (2004) is applied to both groups of countries covering the 1980-2008 period. Once the influence of time-invariant factors is accounted for, the results show a small but significant positive relationship between openness and GDP per worker in both regions. However, the magnitude of the effect is lower in the case of LA. These findings are in line with recent empirical literature.

In terms of economic policy conclusions we can state that the Asian approach to trade openness has been more successful than the LA one. The Asian approach seems to have been more prone to the promotion of export trade and productivity than the LA approach. This is not to say that in LA export promotion was totally absent but it was less successful than in Asia. A possible explanatory reason is that the objective of LA trade reform was to create a neutral trade regime and to let markets work freely. In the LA region production was determined by comparative advantage but as the real exchange rate appreciated this led to a slim import substitution industry and a rather curbed export industry.[23] Besides, as exports in LA are still based on natural resources (mining, wood, fish and crustaceans) and agriculture (soya, maize, wheat, rape seed, meat, fruit and vegetables), the existing production structure might become a threat to productivity in the long run.[24]

In Asia, in contrast, the manufacturing sector is quite strong and, until recently, it has been mostly in the manufacturing sector where modernization and innovation were daily challenges to competitiveness. Nowadays, the service sector, and in particular information technology related services, are also increasingly important activities in those economies. All in all in terms of income generation, the export promoting industrialization process (EPI)

[23] The real exchange rate appreciated primarily in the 1980 to 1999 period in Latin America. In Asia appreciations of the real exchange rate were present in this period as well but the import substitution industry and the export industry were supported by government intervention which reduced the anti-industry bias.

[24] Only Mexico and Brazil were able to produce manufactures to a significant extent.

followed by East and South East countries has proved to be a better strategy than the LA trade liberalization strategy.

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FIGURES

Figure 1. Evolution over time of openness in Asian countries

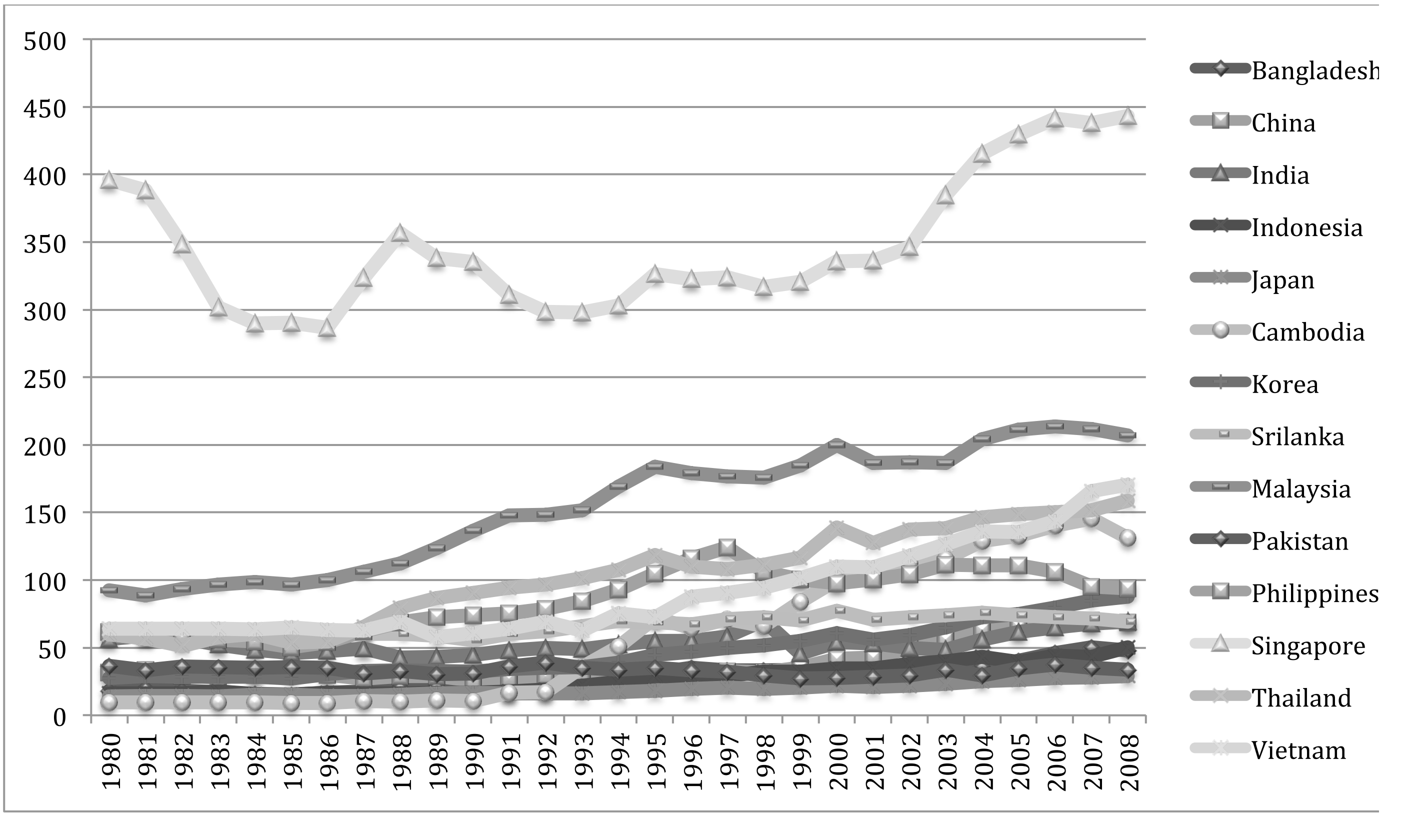
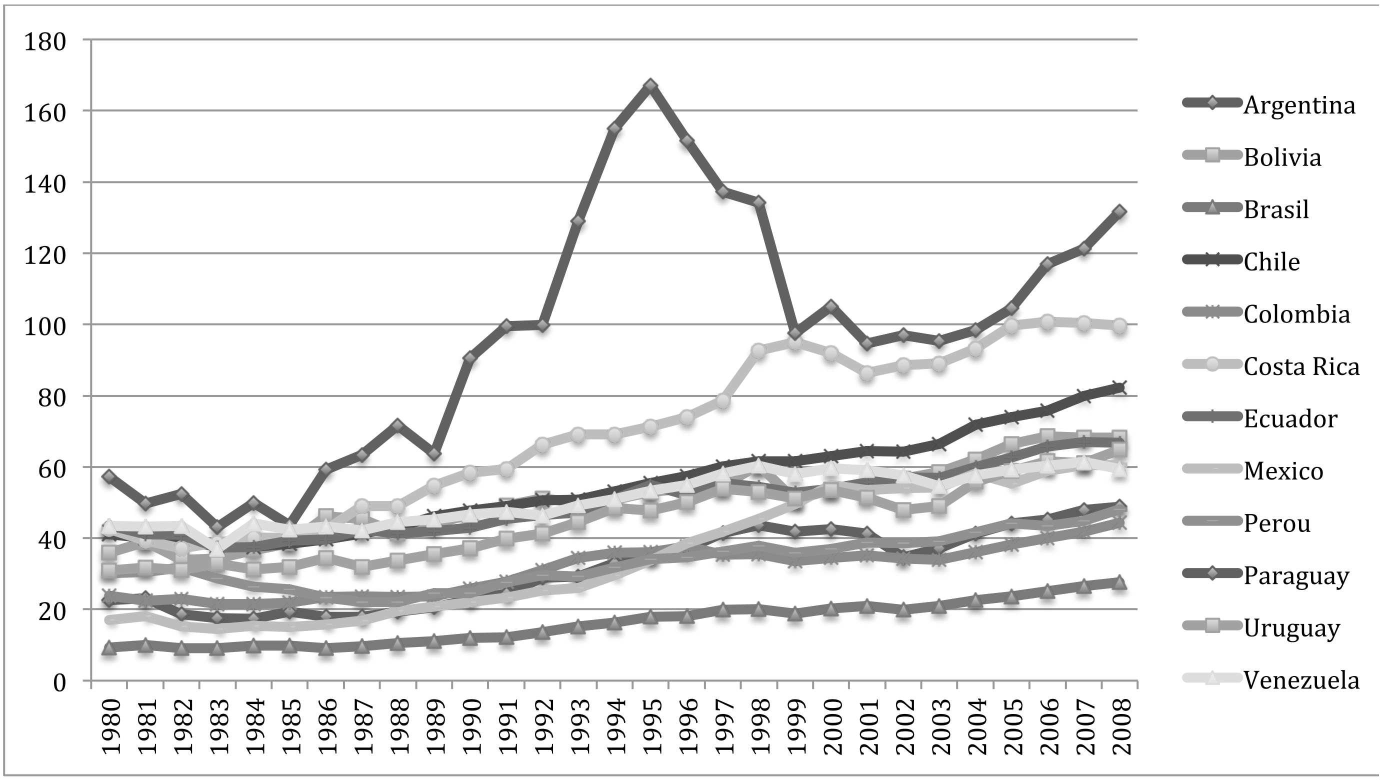


Figure 2. Evolution over time of openness in LA countries



TABLES

Table 1
Pesaran (2004) CD dependence test

Variable	CD dependence test
Asia	
$lprod_{it}$	73.07***
$lopen_{it}$	38.82***
$lpop_{it}$	75.21***
Latin America	
$lprod_{it}$	63.21***
$lopen_{it}$	57.07***
$lpop_{it}$	63.22***

*** denotes rejection at 1% level.

Table 2
Panel Data Statistics based on Approximate Common Factor Models
Bai and Ng (2004) statistics

Panel A: Variables defined for Asia

Bai and Ng (2004b) statistics

	$lprod_{it}$		$lopen_{it}$		$lpop_{it}$	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	--1.896	0.028	0.038	0.515	-7.089	0.00
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-23.33	6	-24.64	6	-27.34	6
MQ test (non-parametric)	-26.97	6	-27.08	6	-27.88	6

Panel B: Variables defined for Latin America

Bai and Ng (2004b) statistics

	$lprod_{it}$		$lopen_{it}$		$lpop_{it}$	
	Test	p-value	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	1.052	0.853	0.017	0.507	-7.357	0.00
	Test	$\hat{r}_1$	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-24.239	6	-24.689	6	-25.83	6
MQ test (non-parametric)	-24.117	6	-26.512	6	-26.96	6

Table 3: Banerjee and Carrion (2010) BC cointegration tests

	Asian countries *f(flopen$_{it}$, lpop$_{it}$)*			Latin American countries *f(flopen$_{it}$, lpop$_{it}$)*		
Model	Z_j^*	AIC	BIC	Z_j^*	AIC	BIC
1	-18.54	-8.22	-7.66	-18.16	-7.48	-6.95
2	-66.86	-8.85	-8.16	-63.17	-7.84	-7.18
3	**-25.95**	**-9.16**	**-8.33**	**-40.68**	**-8.35**	**-7.55**
4	-33.21	-8.74	-7.91	-26.42	-8.17	-7.38
5	-23.63	-8.79	-7.82	-23.67	-8.30	-7.37
6	-28.20	-9.10	-8.00	**-16.92**	**-9.07**	**-8.01**

Table 4
Estimation of the long-run parameters 1980-2008
Asia

Model with a structural break in the intercept and the trend

Variables	LSDV	Bai FM	CupFM	CupBC
t=17 (1996), 1 common factor bandwidth parameter 0.13 (Silverman's rule of thumb)				
$lopen_{it}$	**0.16**	**0.15**	**0.15**	**0.15**
	(6.14)	(5.86)	(5.80)	(5.84)
$lpop_{it}$	**-0.73**	**-0.74**	**-0.76**	**-0.74**
	(-5.89)	(-5.80)	(-5.86)	(-5.80)

Note: LSDV denotes the Least Squares Dummy Variable estimator; Bai FM denotes Bai's Fully Modified estimator; CupFM denotes the Constantly Updated Fully Modified estimator and CupBC denotes the Constantly Updated Banerjee-Carrión estimator. t denotes the year in which an structural break appears.

Table 5
Estimation of the long-run parameters 1980-2008
Latin America

Model with a structural break in the intercept and the trend

Variables	LSDV	Bai FM	CupFM	CupBC
t=6 (1985), 1 common factor bandwidth parameter 0.08 (Silverman's rule of thumb)				
$lopen_{it}$	0.007	**0.07**	**0.08**	**0.08**
	(0.18)	(1.87)	(2.28)	(2.27)
$lpop_{it}$	**-1.25**	**-1.37**	**-1.38**	**-1.38**
	(-9.90)	(-12.40)	(-12.53)	(-12.53)

Model with a structural break in the intercept,
the trend and the cointegration vector

Variables	LSDV	Bai FM	CupFM	CupBC
t=18 (1997), 1 common factor bandwidth parameter 0.08 (Silverman's rule of thumb)				
$lopen_{it}$	**0.11**	**0.09**	**0.098**	**0.11**
	(2.62)	(2.36)	(2.41)	(2.79)
$lpop_{it}$	**-1.40**	**-1.44**	**-1.48**	**-1.52**
	(-9.73)	(-11.32)	(-11.53)	(-11.66)
$lopen2_{it}$	0.08	**0.09**	**0.10**	**0.10**
	(1.57)	(1.85)	(2.07)	(2.13)
$lpop2_{it}$	-0.02	-0.02	0.005	**0.03**
	(-1.62)	(-1.37)	(0.30)	(1.65)

Note: LSDV denotes the Least Squares Dummy Variable estimator; Bai FM denotes Bai's Fully Modified estimator; CUP_FM denotes the Constantly Updated Fully Modified estimator and CUP_BC denotes the Constantly Updated Banerjee-Carrión estimator. t denotes the year in which an structural break appears. Lopen and lopen2 denote the estimated coefficients for real openness before and after the structural break, the same applies to population.

APPENDIX

Table A.1. Sectoral structure of production activities across countries

Year 2000	GDP (Bill. US$)	Agriculture (%)	Industry (%)	Manufacturing (%)	Services (%)
LA					
Argentina	344.3	4	24	15	72
Bolivia	8.4	15	30	15	55
Brazil	644.7	6	28	17	67
Chile	79.3	6	32	17	62
Colombia	99.9	9	29	15	62
Costa Rica	15.9	9	32	25	58
Ecuador	18.3	16	36	19	48
Mexico	683.6	4	35	20	62
Paraguay	8.2	16	36	..	49
Peru	50.7	9	32	17	59
Uruguay	22.8	7	25	14	69
Venezuela	117.1	4	50	20	46
Average		8,75	32,417	17,636	59,083
ASIA					
Bangladesh	47.1	26	25	15	49
Cambodia	3.7	38	23	17	39
China	1,198.5	15	46	32	39
India	476.6	23	26	15	51
Indonesia	165.0	16	46	28	38
Japan	4,731.2	2	31	21	67
Korea, Rep.	561.6	4	38	29	58
Malaysia	93.8	9	48	31	43
Pakistan	74.0	26	23	15	51
Philippines	81.0	14	34	24	52
Singapore	95.8	0	35	28	65
Sri Lanka	16.3	20	27	17	53
Thailand	122.7	9	42	34	49
Vietnam	33.6	23	34	17	43
Average		16,07	34,14	23,07	49,79

Source: World Development Indicators 2014. Industry corresponds to ISIC divisions 10-45 and includes manufacturing (ISIC 15-45). Manufacturing excludes mining and related activities.

Table A2: Banerjee and Carrion (2010) BC cointegration tests
Asian sample excluding Japan

	Asian countries $f(flopen_{it}, lpop_{it})$		
Model	Z_j^*	AIC	BIC
1	-18.89	-8.53	-7.99
2	-25.44	-8.79	-8.11
3	**-51.50**	**-9.09**	**-8.27**
4	-45.90	-8.73	-7.92
5	-22.68	-8.74	-7.79
6	-19.82	-8.92	-7.84

Table A.3
Estimation of the long-run parameters 1980-2008
Asia excluding Japan

Model with a structural break in the intercept and the trend				
Variables	**LSDV**	**Bai FM**	**CupFM**	**CupBC**
t=22 (2001), 1 common factor bandwidth parameter 0.12 (Silverman's rule of thumb)				
$lopen_{it}$	**0.54**	**0.54**	**0.57**	**0.57**
	(8.37)	(10.86)	(11.35)	(11.37)
$lpop_{it}$	**0.30**	**0.31**	**0.31**	**0.31**
	(6.53)	(8.02)	(8.11)	(8.13)

Note: LSDV denotes the Least Squares Dummy Variable estimator; Bai FM denotes Bai's Fully Modified estimator; CupFM denotes the Constantly Updated Fully Modified estimator and CupBC denotes the Constantly Updated Banerjee-Carrión estimator. t denotes the year in which an structural break appears.