Ibero-Amerika Institut für Wirtschaftsforschung
Instituto Ibero-Americano de Investigaciones Económicas
Ibero-America Institute for Economic Research
(IAI)

Georg-August-Universität Göttingen
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Diskussionsbeiträge · Documentos de Trabajo · Discussion Papers

Nr. 226

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September 2013

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TRADE OPENNESS AND INCOME:

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08.09.2013

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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). While the topic even seems to have been over-studied, robust empirical findings are still absent and the issue of how to tackle reverse causation between trade and income (variously taken as levels or growth of GDP or productivity) and cross-section dependence of the countries under study is still unresolved. By utilizing panel cointegration techniques we aim at addressing both problems in one go.

Although there is an extensive empirical literature on the trade-income link we think it is still worthwhile to re-examine this well-documented relationship (Edwards 1992, 1993, 1994 and 1998; Rodríguez and Rodrik, 2001; Baldwin, 2003; Dollar and Kraay, 2003; Lee et al., 2004 and Singh, 2010). First of all, the issue whether trade openness promotes development in terms of income generation has not been settled yet in empirical studies. More recently an interesting finding emerged from studying the time series properties of trade openness and growth of GDP per worker: a) trade openness being a non-stationary variable [I(1)] and growth being stationary [I(0)] cannot be related in a systematic way as these variables cannot be cointegrated; b) trade openness being I(1) and GDP per worker being I(1), in contrast, could be related in a systematic way and over longer periods of time. This can be verified by means of so-called cointegration tests.

Second, the use of real trade openness[1] 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. Nonetheless, it is generally known that the Latin American and the Asian region followed quite different trade regimes over the last 30 years (Narula, 2002). While Latin America

[1] The reasons why real openness should be used instead of nominal openness are detailed in Acalá and Ciccone (2004).

concentrated on trade liberalization, Asia followed a more pragmatic trade policy with much more government intervention. After the severe Latin American debt crisis in 1982, the Latin American countries were forced to follow structural adjustment programs that were composed of trade liberalization, macroeconomic stabilization, deregulation and privatization. Around 1985 basically all Latin American countries had moved to a strategy of (unilateral)[2] trade liberalization implying the abolition of non-tariff and tariff trade barriers. Import tariffs were drastically reduced and export taxes were slashed[3]. The idea was to stop the anti-export 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). In this way, the import substitution industry should become competitive and the sometimes heavily taxed export industry be put into a position to compete on world markets as well. In Asia, in contrast, trade policy did not follow the free market approach which was propagated in Latin America. The import substitution industry was deliberately protected to install first a coal and steel industry. Later on, 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. But government intervention not only played a role in strengthening the import substitution industry but also in 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).

From an economic history perspective, we could imagine a structural break in the mid-1980s in Latin America (due to the structural adjustment programs) and in the late 1990s Asia (due

[2] 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.

[3] 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.

to the burst of a housing bubble and a current account crisis in 1997 which triggered harsh economic reforms).

Acting under different trade regimes, Latin America and Asia had quite similar exchange rate regimes. In most Latin American countries the 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.

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 Latin American 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 of all right hand side variables but also for the autocorrelation of the error terms. Morover, we can be sure not to have an omitted variable problem, if we find the series to be cointegrated. 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 Latin America 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 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 comparative advantage where each country maximizes its welfare by concentrating on those activities in which it is most economically efficient[4]. Generally, according to the neoclassical theory, an opening up towards 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[5]. Finally, although the impact of institutions has not yet been incorporated into economic growth theory, it is widely recognized that without basic institutions[6], 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.

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

[5] Scale effects are derived from the closer integration of a country into the world market, while allocative 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.

[6] Law and order, well-defined property rights and impartially-enforced contracts, economic activity will be constrained. If these XXXX something may be missing???

Krueger (1978) and Bhagwati (1978) started off the debate on whether foreign trade regimes matter for growth. Both 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-country studies was intensified in the 1990s.[7] 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[8] (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,

[7] Barro (1991) established the cross-section growth literature that did not include trade openness at that time.
[8] Which reflects the difference between the official exchange rate and the exchange rate in the black market.

mentalities and so forth. Second, most of the studies paid insufficient tribute to the issue of endogeneity[9] due to reverse causation between the right hand side variables (savings, human capital accumulation, population growth, and openness) and growth.

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). In the infant stage these models were static and pooled OLS (i.e. a common intercept for all countries) was the most favored estimation technique. In a later stage, mostly fixed effects (within estimation) were used to capture time-invariant country characteristics and allowed to control 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). Two types of GMM estimation[10] became popular in later years when the regression equation models were set up as dynamic panel data models. First, difference GMM, in which the equation was run in first differences thus eliminating the fixed effects, became fashionable (Arellano and Bond, 1991). The endogeneity of the lagged dependent variable was tackled by replacing this variable (in levels) by its lagged first differences (or levels). Second, system GMM, which runs the equation in first differences (with the lagged variables in levels as IVs) *and* in levels (with the lagged first differences as IVs), was the recommended option if the series were persistent (non-stationary). This was the approach proposed by Blundell and Bond (2000). 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)

[9] Hsiao (1987) wrote an influential paper on the causality and exogeneity between exports and economic growth.
[10] Which goes back to Hansen (1982).

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 relationship between real openness and labor productivity from the 1990s onwards[11].

A more recent criticism applied to the studies on the trade-income link came from the 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, time-invariant variables are not enough and a proper treatment of the time dimension is needed. 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 play a very important role and in which emphasis is put on whether the series are co-integrated, i.e. move systematically together or apart over a long period of time. Our study falls into this latter category and is not concerned with random co-movements between trade openness and growth that might happen for short periods of time.

These latter studies work with a reduced number of variables, usually two (namely, per capita-income and trade openness). They claim that if co-integration is found between the two variables, the residual is stationary (does not show systematic upward or downward

[11] System-GMM dynamic panel estimation methods are used to address potential biases associated with cross-section estimations such as small sample bias, omitted variable problems and endogeneity of explanatory variables.

movements) and therefore, one can conclude that omitted variables do not destroy the existing co-integrating relationship between trade openness and income per worker. According to Herzer (2013) this justifies a reduced form model 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[12] are added to the bivariate regression. Studies in this line of research are the ones by 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 Dreher 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 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 non-linearities or discontinuities in the opennes-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 viceversa. In fact, the existence of discontinuities can be assimilated to a local aproximation of a non-linear relationship. The lack of control for

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

structural breaks in the series may be reflected in the parameters of the estimated models that, when used for inference or forecasting, can induce 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 abovementioned literature in two different respects. First, we improve the specification adding to the basic bivariate model an extra variable with a time dimension: population. 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 time invariant variables[13], 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 15 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

[13] 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.

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[14]. 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, 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 (2010). 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 to note that the countries analyzed in this paper belong to two rather homogeneous samples (Asia and Latin America). This fact makes

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

the assumption of a homogenous break for each regional bloc rather suitable.

Finally, we employ Bai et al. (2009) 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 Alcalá and Ciccone (2004) model, where instead of including geographical or institutional variables, we control for them trough 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 Latin American 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)[15]. The

[15] The sample spans from 1980 to 2008 and includes two groups of countries: Asia and Latin America. The Asian 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), olivia (BOL), Brazil (BRA), Chile (CHL), Colombia (COL), Costa Rica (CRI), Equator (ECU), Mexico (MEX), Paraguay (PRY), Peru (PER), Uruguay (URY) and Venezuela (VEN).

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[16] over time periods $t= 1, 2,\ldots,T$ and countries $i=1, 2,\ldots,N$; $lopen_{it}$ denotes a measure of real openness[17] 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 (which does not become relevant in the case of cointegration) by including direct proxies for geography, and institutions, since it can be 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 Rodrick (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

[16] Labor Productivity is defined as $\left(\frac{PPPGDP_{it}}{Work_{it}}\right)$

[17] 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).

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 yields 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). 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. 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

There are two important aspects that should be taken into account before estimating the cointegrating relationship. First, it is highly probable that the series are interrelated among them, since the countries in the sample are members of highly integrated areas, Latin America and Asia. 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.

A second important point is to test for the existence of unit roots in the data. 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 $Y_{i,t}$, as follows:

$$Y_{i,t} = D_{i,t} + F_t' \pi_i + e_{i,t}, \qquad (2)$$

with $t = 1, \ldots, T$, $i = 1, \ldots, N$, where $D_{i,t}$ denotes the deterministic part of the model — either a constant or a linear time trend, F_t is a (r x1)-vector that includes the common factors that are present in the panel, and $e_{i,t}$ is the idiosyncratic disturbance term, which is assumed to be cross-section independent. Unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model. 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 BIC information criterion 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 Latin America 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 Latin American 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 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 (2010) 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 worth noticing that inference concerning the presence of cointegration can be affected by misspecification if the existence of breaks is ignored.

Banerjee and Carrion-i-Silvestre define a *(m x 1)* vector of non-stationary stochastic process, $Y_{i,t} = (y_{i,t}, x_{i,t}^{'})$ whose elements are individually I(1) with the following Data Generating Process:

$$y_{i,t} = D_{i,t} + x_{i,t}^{'}\delta_{i,t} + u_{i,t} \tag{3}$$

The general functional form for the deterministic term $D_{i,t}$ is given by:

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{m_i} \theta_{i,j} DU_{i,j,t} + \sum_{j=1}^{m_i} \gamma_{i,j} DT_{i,j,t}, \tag{4}$$

where $DU_{i,j,t}$ =1 and $DT_{i,j,t}$ = (t - $T^b_{i,t}$) for $t > T^b_{i,t}$ and 0 otherwise, $T^b_{i,t} = \lambda^b_{i,j} T$ denotes the timing of the *j*-th break, $j = 1,\ldots, m_i$, for the i-th unit, $I = 1,\ldots, N$, $\lambda^b_{i,j} T \in \Lambda$, being Λ a closed subset of (0,1). The cointegrating vector is a function of time so that

$$\delta_{i,t} = \begin{cases} \delta_{i,1} T^c_{i,0} < t \le T^c_{i,1} \\ \delta_{i,2} T^c_{i,1} < t \le T^c_{i,2} \\ \cdots\cdots \\ \delta_{i,j} T^c_{i,j-1} < t \le T^c_{i,j} \\ \cdots\cdots \\ \delta_{i,n_i+1} T^c_{i,n_i} < t \le T^c_{i,n_i+1} \end{cases} \tag{5}$$

with $T^C_{i,0} = 0$ and $T^C_{i,n_i+1} = T$, where $T^C_{i,j} = \lambda^C_{i,j} T$ denoting the *j*-th time of the break, $j = 1,\ldots,n_i$, for the *i*-th unit, $i = 1,\ldots,N$, for the $i-$th unit, $i = 1,\ldots,N$, $\lambda^C_{i,j} \in \Lambda$.

Banerjee and Carrion-i-Silvestre (2010) 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 (2010) recover the idiosyncratic disturbance terms ($\tilde{e}_{i,t}$) 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 $t^{j}_{\tilde{e}_i}(\lambda_i)$, $j = c, \tau, \gamma$. In the models without a time trend (Models 1 and 6) j=c. In those with a stable linear time trend (Models 2, 3 and 7) j=τ and, finally, j=γ in the models with a changing trend (Models 4, 5 and 8).

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 Latin American countries. Using the Bayes information criterion (BIC) [18] we choose Model 5 (in Table 3) in the case of the Asian countries, that is, the one that contains a constant and a trend that is affected by a structural break. For Latin America, we select both, Model 8 (Table 3), including a structural break that affects not only the deterministic components but also the cointegrating vector and, again, Model 5 (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 5) or 1997 (according to model 8) for Latin American 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

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

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 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 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:

$$y_{it} = x_{it}^{'}\beta + e_{it} \tag{6}$$

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

$$x_{it} = x_{it-1} + v_{it} \tag{7}$$

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:

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

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}:

$$e_{it} = \lambda_i' F_t + u_{it} \tag{9}$$

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,

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

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 5 and 8 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 8 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)[19]. 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 Latin American countries, for comparative purposes we have estimated the same specification as for Asian countries, with breaks in the intercept and the slope (Model 5) together with Model 8. In Model 5, 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 8. 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 8 for Latin American countries and focus on Model 5.

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 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.

[19] 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.

Concerning the deterministic components of the estimated model, we find a structural break in 1984 for Latin America and in 1999 for Asia. 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 Latin America, 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. Latin American countries had a neutral trade regime, whereas in Asia the government played an active role in shaping competitive advantage. Asian industries (both import substitution and export industries) were promoted by deliberate government intervention.

As for the effect of population on productivity, it is negative both in Asia and in Latin America. 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 Latin America this reduction is almost twice (1.38%). This result is consistent with the theory.

All in all, the export promoting industralization process (EPI) followed by East and South East countries has proved to be a better strategy than Latin American import substitution industrialization (ISI).

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, about the relationship between economic performance and trade openness. Despite voluminous work in this area, the findings are far from being conclusive. Asian and Latin American 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 influenced of time-invariant factors are 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 Latin America. 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 Latin American one. The Asian approach seems to have been more prone to the promotion of export trade and productivity than the Latin American approach. This is not to say that in Latin America export promotion was totally absent but it was less successful as in Asia as the objective of Latin American trade reform was to create a neutral trade regime and to let markets work freely. In the Latin American 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.[20] Besides, as exports in Latin America are primarily 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.[21]

[20] 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.

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

In Asia, in contrast, the manufacturing sector is quite strong and after all, it is in the manufacturing sector where modernization and innovation are daily challenges to competitiveness.

Acknowledgements

The authors are indebted to J. LL. Carrión-i-Silvestre and Ch. Kao for providing them with the Gauss codes to implement some of the tests used in the paper. The usual disclaimer applies. M. Camarero and C. Tamarit acknowledge the funding from *MICINN* (project ECO2011-30260-C03-01). Both are members of INTECO research group funded by *Generalitat Valenciana,* PROMETEO/2009/098 project. F. Nowak-Lehman D. and I- Martínez-Zarzoso acknowledge the funding from *MICINN* (project ECO2010-15863)

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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	-5.47	-8.31	-7.90	-9.66	-7.44	-7.04
2	-13.48	-8.94	-8.39	-12.53	-7.76	-7.23
3	-18.54	-8.22	-7.66	-18.16	-7.48	-6.95
4	-66.86	-8.85	-8.16	-63.17	-7.84	-7.18
5	**-25.95**	**-9.16**	**-8.33**	**-40.68**	**-8.35**	**-7.55**
6	-33.21	-8.74	-7.91	-26.42	-8.17	-7.38
7	-23.63	-8.79	-7.82	-23.67	-8.30	-7.37
8	-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.