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Financial Development and Economic Activity in Advanced and Developing Open Economies: Evidence from Panel Cointegration

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Abstract

This study considers the effects of financial development on output in a panel cointegration framework, focusing on the implications of trade and financial openness. Our analysis indicates that after controlling for cross-sectional dependence the typical relationship between finance and output does not hold in the long-run. This relationship, however, is re-established once we account for economic openness. While trade openness emerges as more important for developing countries financial openness is more important for advanced economies. In the long-run, causality runs from financial development to output in the advanced economies, while in developing economies causality is bidirectional. There is no short-run causality between financial development and output, however.

1. Introduction

The role of the financial system in promoting economic growth has been highlighted more than a century ago (Schumpeter, 1911) and its importance has been stressed in various contexts over time. The work of McKinnon (1973) and Shaw (1973) highlight the role of financial liberalization in bringing about financial development, which in turn can spur economic growth. The theoretical models of endogenous growth developed in the 1980s incorporate explicitly financial institutions[1] and analyze the channels that allow financial intermediaries to affect growth. Alongside the theoretical contributions, an extensive empirical literature emerged, which on balance supports the existence of a finance-growth nexus.

A number of authors have emphasized the interconnectedness across countries and its implications for financial development and growth. Rajan and Zingales (2003) provide a theoretical rationale linking international trade and finance to financial development as intensified competition and investment opportunities incentivize domestic firms to draw on external finance. While some research emphasizes the importance of trade openness (e.g., Alcala and Ciccone, 2004; Lee *et al.*, 2004; Dollar and Kraay, 2003) others focus on the openness (or lack of restrictions) in capital movements (e.g., Edison *et al.*, 2002; McKenzie, 2001). The more open an economy is the more likely is to engage in policies that promote financial development. In other words countries that encourage financial deepening are likely to also opt for more openness. To our knowledge, however, no panel evidence exists from nonstationary data considering the link between financial development and growth when the openness of the economy is taken into consideration.

This paper uses a recently developed panel cointegration framework that takes into account cross sectional dependencies to examine the role of financial deepening in economic activity in open economies, considering explicitly trade and financial openness. The empirical literature on financial development and development focuses on stationary panels (e.g., Levine *et al.* 2000) as well as on time series methods (e.g., Arestis *et al.*, 2010), including cointegration analysis (e.g., Christiopoulos and Tsionas, 2004). Nevertheless, very limited evidence exists on nonstationary panels. Moreover, the existing panel cointegration studies that consider the nexus between finance and economic activity typically do not account for cross-sectional dependence in the cointegrating relationship

[1] See for instance, Greenwood and Jovanovic (1990) and Bencivenga and Smith (1991).

and for the possibility of cointegration across cross-sectional units when testing for stationarity.

The findings and contribution of this paper is threefold. First, we show that once cross-section dependence is taken into account there is no long-run relationship between financial development and output. Second, when financial and trade openness are considered then a long-run relationship between financial deepening and output emerges. We find that financial development and trade openness are more important for the developing economies while financial development along with financial openness is more important for the advanced economies. Third, causality analysis reveals that the relationship between financial development and output is a long-run phenomenon and not a short-run one. The direction of causality runs from financial development to output for the advanced economies and is bidirectional for the emerging economies.

The rest of the paper is structured as follows. Section 2 reviews the literature, Section 3 presents the data and the methodology, Section 4 provides the empirical results, and Section 5 concludes.

2. Background

An abundance of evidence has been produced to gauge the effects of financial development on economic growth. The related literature uses various econometric approaches and covers a wide variety of countries and time period spans. King and Levine (1993) construct four measures of financial development for 80 countries and analyze cross-country data over the period 1960-1989. Their findings reveal a positive and statistically significant effect of the financial variables on real per capita GDP growth. The subsequent studies of Demirguc-Kunt and Maksimovic (1998) and Levine and Zervos (1998) provide additional evidence for this positive effect of financial development on economic activity.

Early studies have been criticized for not taking account potential endogeneity. To remedy this methods based on instrumental variables have been employed. Harris (1997) within a 2SLS framework, finds that the beneficial effects of the stock market activity are limited only to the developed economies. Furthermore, Levine (1998, 1999, 2002) and McCaig and Stengos (2005), employ GMM and find that growth is positively associated with financial development proxies. Similar results are provided by Beck *et al.* (2000), Benhabib and Spiegel (2000), Levine *et al.* (2000) and Henry (2000).

A number of studies focused on the distinction between bank-based and market-based financial systems and their corresponding influence on growth. Beck and Levine (2004) and Ndikumana (2005) suggest that both forms of financial intermediation play a significant role. Moreover, Rousseau and Wachtel's (2000) empirical investigation demonstrates the increasing influence of stock markets on economic activity for both developed and developing economies; a well functioning stock market, in terms of sufficient liquidity, affects positively real output.

Rioja and Valev (2004) argue that the level of financial development determines the significance of the effects. Dividing the examined countries into three groups according to the level of financial development, they find that in economies with less developed financial system the effect of financial development on growth is ambiguous. On the contrary, the effect is positive but small in economies with well-developed financial system. However, for economies that belong to the intermediate group the effect is found to be positive and relatively larger. Deidda and Fattouh (2002) consider a threshold type of nonlinearity and find that a significant relationship between financial development and growth exists only above a specific threshold level of the initial per capita income. Economies with high initial income are characterized by a positive finance-growth link, whereas for countries with low initial income there is no significant link. Henderson *et al.* (2013) also produce evidence of strong nonlinearities in the relationship between financial development and growth. Moreover they find that this effect is very limited in low-income countries.

A separate spate of papers considers the relationship between growth and financial openness.[2] Quinn (1997) measures capital account openness by constructing an index based on cross-border financial transactions reported by IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER). Quinn's measure, like several other indices based on AREAER, represent a "de jure" index.[3] His findings point towards a positive association between growth and financial liberalization. On the other hand, Rodrik's (1998) results, also based also on a de jure financial openness variable, suggest that such relationship does not exist. McKenzie's (2001) results corroborate

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[2] The terms financial liberalization or integration are used in an alternated way.

[3] De jure indices contained information about the financial openness that is based on official declarations and statements about capital movements. However, as we will discuss in the next section, de jure measures may not always reflect the process of financial openness in an appropriate and systematic way. For this reason "de facto" indices have been constructed which are based on actual flows of capital movements.

those of Rodrik. Those studies that take into account endogeneity and reverse causality, however, provide mixed evidence.[4] Eichengreen and Leblang (2003), using a de jure measurement of financial integration, support the view that in periods of financial stability more open economies tend to grow faster. Edison *et al.* (2002) employ both de jure and de facto indices and find that there is no robust link between financial openness and growth.

Other authors stress the special aspects of the liberalization process. For instance, Klein (2005) emphasizes the importance of institutions, while Chanda (2005) finds growth gains for more ethnically homogeneous countries. Quinn and Toyoda (2008) suggest that capital account openness as well as equity market liberalization contribute significantly to growth. Bussiere and Fratzscher (2008) provide evidence that the growth benefits hold only for the short-run. The quality of domestic institutions and the size of FDI inflows determine whether the positive growth effects will continue in the medium and long-run term.[5]

A relatively limited part of the empirical literature investigates the connection between growth and trade openness. In their survey, Rodriguez and Rodrik (2000) casts doubts about the beneficial effects of trade on growth found in previous studies. Subsequent research supports the positive effects of trade; see Irwin and Tervio (2002), Dollar and Kraay (2003), Alcala and Ciccone (2004) and Lee *et al.* (2004). Baltagi et. al. (2009) examine if trade and financial openness can explain the pace in financial development, and its variation across countries, finding that both types of openness affect banking sector development.

Another strand of the literature adopts a time series approach and examines the long-run relationship between financial development and economic activity (real GDP) (instead of growth) using cointegration analysis.[6] Time series analyses typically employ the Johansen cointegration methodology and provide evidence in favor of a long-run relationship between real GDP and financial development in both developed and developing economies. Such studies include Rousseau and Wachtel (1998), Luintel and Khan (1999), Arestis et al. (2001) and Rousseau and Vuthipadadorn (2005). Panel cointegration methods arrive to similar conclusions (Christopoulos and Tsionas, 2004;

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[4] The most frequently used method of this branch of literature is dynamic panel GMM.

[5] A part of the related literature has focused on the relations between FDIs (as a specific aspect of financial openness) and growth (see Alfaro *et al.* (2004) and Ang (2009)).

[6] Growth rates are stationary series and as a result cointegration cannot be employed.

Apergis *et al.*, 2007; Neusser and Kugler, 1998). However, these studies do not account for cross-sectional dependence in the cointegrating relationship and the possibility of cointegration across cross-sectional units when testing for stationarity. Moreover, to the best of our knowledge the role of economic openness has not been considered in a panel framework.

3. Data and Methodology

3.1 Models and Data

The empirical literature on the finance-growth nexus typically computes averages for several sub-periods and, assuming that all variables are stationary, employs dynamic panel GMM techniques to explore the relationship between economic growth and financial development. We follow a different perspective and examine the existence of a long-run relationship linking output [economic development] with financial development and the openness of the economy, using cointegration analysis

Given the observed heterogeneities between developed and less-developed economies, we split our country sample into two groups; the first one includes 20 advanced countries while the second one contains 17 emerging economies[7]. The data are annual and the time span covers the period 1970-2007.

We start with the investigation of the relationship between output and financial development as the benchmark case examined by the related literature. The possible cointegration relationship in the simple two-variable equation is:

$$y_{i,t} = c_i + a_{1i} fd_{i,t} + u_{i,t} \tag{1}$$

where y is the logarithm of real per capita GDP converted to US dollars (base year 2000). *fd* is a proxy of financial development as measured by the domestic credit provided by financial institutions to the private sector as a percentage of GDP. Both variables are taken from the World Development Indicators (WDI) constructed by the World Bank.

However, this framework can be restrictive in the sense that additional variables highlighted in the literature need to be accounted for. The financial development proxy used here measures the degree to which firms have access to finance from domestic

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[7] Particularly, the examined developed countries are: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Malta, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, UK and US. The emerging economies group is composed of Argentina, Brazil, Chile, Colombia, Egypt, India, Iran, Korea, Malaysia, Mexico, Pakistan, Peru, Philippines, Singapore, South Africa, Thailand and Venezuela.

banks and other financial institutions. This variable provides information about domestic financial conditions. However, the private sector could be financed from foreign sources. The greater the number of alternative external finance opportunities in an economy, the more open this economy is. These external financial sources can take a various forms. To take their combined influence into account, we also consider a composite variable aimed at capturing the degree of financial openness. In addition, we also include another variable that purports to capture another aspect of the degree of a country's openness, the trade openness, commonly measured in empirical work as (imports + exports)/GDP. Table 1 provides a summary of the data.

Table 1 here

Given the discussion above, the testing equation can now be written as:

$$y_{i,t} = c_i + a_{1i}fd_{i,t} + \beta_{1i}fo_{i,t} + \gamma_{1i}to_{i,t} + v_{i,t} \qquad (2)$$

where *fo* and *to* stand for the logarithm of financial and trade openness respectively. For financial openness we use the extended database of Lane and Milesi-Ferretti (2007) in which the gross assets and liabilities for a large set of countries are reported. We gauge financial openness using the ratio of the stock of total flows of foreign assets and liabilities to GDP. The stock includes the inflows and outflows of Foreign Direct Investments (FDIs), portfolio equity investments as well as the external debt flows. Subsequently, portfolio equity inflows are defined as the foreign investors' purchases of domestically issued equity in a company. Analogously, portfolio equity outflows refer to the domestic investors' purchases of foreign issued equities. Debt inflows are considered the foreigners' purchases of debt issued by corporate or the government as well as foreign borrowing undertaken by domestic banks. In a similar way, debt outflows refer to the domestic investors' purchases of foreign debt and the domestic borrowing by foreign banks.

We employ the "de facto" index since it reflects the entire process of capital openness irrespective of what one country may officially declare, which is measured by the "de jure" indicators. For instance, many Latin American countries have imposed a number of restrictions on capital outflows. However, this did not avert the huge capital

flights of the last two decades.[8] Lastly, trade openness is measured by the volume of exports plus imports to GDP provided by WDI.

3.2 Testing for Unit Roots and Cross-Sectional Dependence

We examine the order of integration of the series by employing three panel unit root tests. We start with the IPS test proposed by Im, Pesaran and Shin (2003) which is commonly used in empirical work. A shortcoming of this test is that it does not take into account the possible cross-sectional dependence among the variables of the panel. Such dependence distorts the inference as the asymptotic analysis is no longer accurate. For this reason we adopt the Pesaran (2004) CD test for dependence. We then proceed to perform the panel unit root tests proposed by Pesaran (2007) and Chang and Song (2009) that take into account cross sectional dependence.

The ADF regression equation can be written as:

$$\Delta y_t = a + \beta y_{t-1} + \sum_{j=0}^{p} \gamma_j \Delta y_{t-j} + \varepsilon_t \tag{3}$$

The IPS test is derived as a simple average of individual ADF unit root tests:

$$t_{IPS} = \frac{\sqrt{N}(\bar{t} - E(t_i))}{Var(t_i)} \tag{4}$$

where N is the number of cross-sectional units, t_i refer to the individual ADF t-statistics and $\bar{t}$ is the corresponding average, i.e., $\bar{t} = N^{-1} \sum_{i=1}^{N} t_i$ and $E(t_i)$ and $Var(t_i)$ are the mean and the variance of $\bar{t}$. This statistic is proved to be normally distributed. However, IPS assume that there are no dependencies among the series. For this reason, we perform the CD test for dependence (Pesaran 2004). The first step consists of estimating the simple ADF equation (3) for each cross-section separately. Secondly, we compute the pair-wise cross-section correlation coefficients of the residuals from the equations (3), $\hat{\rho}_{i,j}$, and compute the simple average of these coefficients across all the $N(N-1)/2$ pairs, $\bar{\bar{\rho}}$, which is equal to:

$$\bar{\bar{\rho}} = \left(\frac{2}{N(N-1)}\right) \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{i,j} \tag{5}$$

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[8] See Kose *et al*. (2009).

Pesaran (2004) shows that the test:

$$CD = \left(\frac{TN(N-1)}{2} \right)^{1/2} \bar{\hat{\rho}} \qquad (6)$$

is normally distributed. If the null of independence is rejected, the inference from IPS is no longer valid.

Pesaran (2007) proposed a test which remedies this drawback incorporating the lag of cross-section mean of $y_{i,t}$, i.e., $\bar{y}_t = N^{-1} \sum_{j=1}^{N} y_{j,t}$ and its differences into the ADF equation and the cross-sectionally augmented Dickey-Fuller (CADF) is given by:

$$\Delta y_{i,t} = a_i + \beta_i y_{i,t-1} + c_i \bar{y}_{t-1} + \sum_{j=0}^{p} d_j \Delta \bar{y}_{t-j} + \varepsilon_{i,t} \qquad (7)$$

Estimating equation (7) for each single cross-section unit, the Pesaran (2007) test statistic can be obtained as the cross-sectional average of the -ADF statistic; the cross-sectional IPS (CIPS):

$$t_{CIPS} = N^{-1} \sum_{i=1}^{N} t_i \qquad (8)$$

The latter does not follow the normal distribution but simulated critical values are available.

Chang and Song (2009) (CS) point towards another possibility that can cause distortions; the presence of cointegration across cross-sectional units. This can be considered as a long-run dependence. The panel unit root test developed by Chang and Song (2009) is given by:

$$\Delta y_{i,t} = \beta_i y_{i,t-1} + \sum_{k=0}^{p_i} f_{i,k} \Delta y_{i,t-k} + \sum_{k=0}^{q_i} g_{i,k} w_{i,t-k} + \varepsilon_{i,t} \qquad (9)$$

where w (covariates) are lagged differences of the remaining cross sections and linear combinations of the lagged levels of all cross sections. Equation (9) is estimated through non-lineal IV. The instruments are generated by a set of "Instrument Generating Functions" (IGFs), $F_i(y_{i,t-1})$. This is an extension of Chang (2002) in which the instruments are generated by a single IGF for all cross-sectional units. Thus, based on the IV estimation of (9) two statistics are proposed: The first one is

$$S = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} t_i \qquad (10)$$

where t_i is the *t*-statistic of β_i. The second statistic is the minimum S and it is actually the minimum t_i, i.e.,

$$S_{\min} = \min_{1 \le i \le N} t_i \tag{11}.$$

The average S tests the null hypothesis that all cross-sectional units contain one unit root against the alternative that all units are stationary. On contrary, $S_{\min}$ tests the same null against the alternative that only a proportion of series are stationary. Here, we compute three alternative versions of these two tests. The first one is based on Chang (2002) where only one IGF is used and there are no covariates. Following CS, we depict these tests as S^C (for the average) and $S^C_{\min}$ (for the minimum). Secondly, a set of IGF is used and again there are no covariates ($S^H, S^H_{\min}$). For the final case, both a set of IGF as well as w covariates are used ($S^A, S^A_{\min}$). It is shown that the average tests are normally distributed, while the critical values for the minimum tests are computed and provided by CS.

3.3 Testing for Cointegration

After the examination of the order of integration, we proceed to the investigation of the long-run relationships. To account for dependence in the context of cointegration, we apply a test developed by Westerlund and Edgerton (2007) and constitutes a modification of the residual-based LM test of McCoskey and Kao (1998). The last show that

$$\sqrt{N}\left(\frac{LM - E(LM)}{Var(LM)}\right) \sim N(0,1) \tag{12}$$

where $LM = \dfrac{1}{NT^2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\omega}_i^{-2}S_{i,t}^2$ with $\hat{\omega}_i$ and $S_{i,t}$ being the estimated long-run variance and the partial sum process of the residuals. This statistic is based on a specific data generating process which can be written as

$$y_{i,t} = a_i + x'_{i,t}b_i + z_{i,t} \tag{13}$$

where the regressors $x_{i,t}$ are assumed to be pure random walk processes and $z_{i,t}$ is the error term which is decomposed into two terms; $z_{i,t} = u_{i,t} + v_{i,t}$. The first term is a stationary process and for the second one, $v_{i,t} = \sum_{j=1}^{t}\eta_{i,j}$ holds with $\eta_{i,j}$ being an

independent and *iid* process with $E(\eta_{i,j}) = 0$ and $Var(\eta_{i,j}) = \sigma_i^2$. The null hypothesis of cointegration against the alternative of no cointegration is reduced to testing whether the variance equals zero, i.e., $H_0 : \sigma_i^2 = 0$ for all *i* against $H_1 : \sigma_i^2 > 0$ for some *i*.

Westerlund and Edgerton (2007) suggest bootstrapping as a way to deal with cross-sectional dependence. This requires the computation of the empirical distribution of this test. The residuals are assumed to admit to an AR(∞) representation. Taking the stationary component $u_{i,t}$ and the first differences of the regressors, $\Delta x_{i,t}^{/}$, which are by definition stationary, they define the vector $w = (u_{i,t}, \Delta x_{i,t}^{/})^{/}$ and the infinite autoregressive representation as:

$$\sum_{j=0}^{\infty} \varphi_{i,j} w_{i,t-j} = e_{i,t} \tag{14}$$

where $e_{i,t}$ is a stationary process. Approximating equation (14) with an autoregressive model of finite order *p*, they employ a sieve bootstrap scheme.[9] At the end of the process, new bootstrap values for the initial variables $y_{i,t}$ and $x_{i,t}$ are produced. Replicating the whole process *N* times and computing each time the **LM** test, the bootstrap distribution is obtained.

4. Empirical Findings

4.1 Order of Integration

Table 2 presents the results for the IPS, CD and PES. the IPS test rejects null that all output and all trade openness variables have a unit root at the 5% significance level for the group of advanced countries and accepts the null in the rest of the cases under consideration..However, the CD test rejects, the null of cross-sectional independence in all cases, implying the inference from IPS is not valid. , The PES test rejects the null of a unit root for all panels only for the financial openness variable of the emerging economies.

Table 2 here

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[9] See also Godfrey (2009).

Table 3 presents the results for the six alternative versions of the panel unit root test proposed by Chang and Song (2009) which takes into account not only cross-sectional dependence but also possible long-run relationship between the variables. For all cases, there is clear evidence that the series are I(1) as the null of unit root cannot be rejected, as all test statistic values are insignificant.

Table 3 here

4.2 Testing for a Long-run Relation

The overall evidence is that all variables are I(1) and we proceed with the estimation of the cointegrating regressions (1) and (2). The results for the former are presented in Table 4 where both asymptotic and bootstrap p-values are reported. The null hypothesis of cointegration is rejected, regardless of the country group. Thus, the long-run link between economic activity and financial development seems to be rather weak when cross sectional dependence is taken into account. This finding is contrary to the results in earlier studies that did not exploit the increased power of combining both the cross-section and the time series dimension and accounting for CD dependence simultaneously.

Table 4 here

We extend the bivariate case by taking into account financial openness and trade openness. Table 5 presents the results for equation (2). The evidence is in favour of a cointegrating relationship. Although the asymptotic p-values indicate the rejection of the null of cointegration, the inference based on bootstrap p-values suggest the opposite (bootstrap tackles CD). These results suggest that the there is a long-run (cointegrating) relationship which, however, cannot be depicted by the bivariate case.

Table 5 here

4.3 Panel Estimation and Causality

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The next step is to quantify the long-run coefficients and establish the direction of causality. Using Fully-modified OLS (FMOLS), the results for equation (2) are presented in Table 6. All right hand side variables are significant at the 1% level of significance. Financial development (and trade openness) is more important for economic development in the emerging economies. On contrary, financial openness is a more important factor for the advanced countries as the corresponding estimates are higher. The latter highlights the need for developing economies to enhance their domestic financial system and accelerate export orientation. For developed economies financial openness emerges as more important reflecting the maturity of their financial system. The estimated values for the 37-countries group (ALL) lie very close to the mean of the corresponding estimates from advanced and emerging groups. Overall trade openness is the economically most significant with a long-run coefficient 33% higher in developing economies.

The coefficients of financial development (and trade openness) are higher in the emerging economies than in the advanced economies. On contrary, financial openness has a higher coefficient in the advanced countries than in the emerging economies. The latter highlights the need for developing economies to enhance their domestic financial system and accelerate export orientation. For developed economies financial openness emerges as more important reflecting the maturity of their financial system.

Table 6 here

So far we have established the long-run relationship between FD and output cannot be reflected in the bivariate case but what about causality? Given that the examined variables are non-stationary but co-integrated, Engle and Granger (1987) showed that an error correction representation can be estimated. Based on the cointegrating equation (2) an ARDL(1,1,1,1) dynamic panel specification can be written as follows:

$$y_{it} = c_i + \lambda_i y_{it-1} + \delta_{i10} fd_{it} + \delta_{i11} fd_{it-1} + \delta_{i20} fo_{it} + \delta_{i21} fo_{it-1} + \delta_{i30} to_{it} + \delta_{i31} to_{it-1} + \varepsilon_{it} \qquad (16)$$

All variables are as previously defined with small letter denoting natural logarithms. The panel error correction reparameterization of the ARDL(1,1,1,1) is

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$$\Delta y_{i,t} = \psi_{1i}(y_{i,t-1} - \theta_{i0} - \theta_{i1}fd_{it} - \theta_{i2}fo_{it} - \theta_{i3}to_{it}) + \delta_{i11}\Delta fd_{it} + \delta_{i21}\Delta fo_{it} + \delta_{i31}\Delta to_{it} + u_{it} \qquad (17)$$

where $\varphi_i = -(1-\lambda_i)$, $\theta_{i0} = \dfrac{c_i}{1-\lambda_i}$, $\theta_{i1} = \dfrac{\delta_{i10}+\delta_{i11}}{1-\lambda_i}$, $\theta_{i2} = \dfrac{\delta_{i20}+\delta_{i21}}{1-\lambda_i}$, $\theta_{i3} = \dfrac{\delta_{i30}+\delta_{i31}}{1-\lambda_i}$.

The first right-hand term stands for the error correction term. Since equation (17) is nonlinear in the parameters, it is estimated through a maximum likelihood proposed by Pesaran *et al.* (1999), the Pooled Mean Group (PMG) estimator. This method allows the intercept, the short-run coefficients and the error variances to differ across the i groups. The only constraint is imposed on the coefficient of the error correction term, ψ_{1i}, which is restricted to be equal across the i groups. The statistical significance of this coefficient is of importance since it provides information about long-run causality. A statistically significant $\hat{\psi}_{1i}$ provides evidence that the financial development leads GDP. So as to test whether GDP leads the financial development we estimate the corresponding error correction model where the financial development is treated as the left-hand side variable;

$$\Delta fd_{i,t} = \psi_{1i}(y_{i,t-1} - \theta_{i0} - \theta_{i1}fd_{it} - \theta_{i2}fo_{it} - \theta_{i3}to_{it}) + \gamma_{i11}\Delta y_{it} + \gamma_{i21}\Delta fo_{it} + \gamma_{i31}\Delta to_{it} + v_{it} \quad (18)$$

A statistically significant $\hat{\psi}_{2i}$ provides evidence that GDP leads financial development. Evidence for short-run causality stems from the estimated coefficients of the first differenced terms: H_0: $\delta_{i11} = 0$ and H_0: $\gamma_{i11} = 0$ in equations (17) and (18), respectively.

As a robustness check, we also compute the Canning and Pedroni (2008) panel test for detecting long-run causality. The latter is based on the error correction models described by equations (17) and (18) estimated for each country separately. The group-mean (GM) test is constructed by the panel average of t-statistics of ψ_i coefficients, i.e., $t = N^{-1}\sum_{i=1}^{N} t_i$ and follows the standard normal distribution under the null of no long-run causal effect.

Firstly, for all groups, there is no evidence of short-run causality from financial development to GDP. On contrary, as far as the long-run causality is concerned, both PMG-L and GM tests support that financial development leads economic activity; all the

p-values are smaller than 5%. However, reverse causality running from economic activity to financial development does not seem to hold. Given that the null of no causality is equivalent to testing the null of weak exogeneity, our results provide evidence that the financial development variable is weakly exogenous (see Enders, 2010). For almost all cases, the PMG-L and GM *p*-values are above 5%. The only exception is the PMG-L test value for the Emerging market group for which there is evidence of bi-directional causality. Overall, the evidence is in favour of unidirectional long-run causality running from financial deepening to economic activity providing support for the Schumpeterian argument. Concerning short run causality, there is no evidence for short-run causality in any direction at the 5% level of significance. However at the 10% level, there seems to be some evidence of unidirectional short run causality running from economic activity to financial deepening in advanced countries.

Table 7 here

5. Concluding Remarks

This study contributes to the literature on the "finance- growth" nexus by employing recently developed panel cointegration techniques and by explicitly considering the role of openness. In particular, we use a test developed by Westerlund and Edgerton (2007), which is based on the residual-based LM test of McCoskey and Kao (1998) and allows accounting for cross-sectional dependence in the cointegrating relationship. We also make the distinction between trade and financial openness.

Our findings suggest that a long-run relationship between financial development and output does not exist when the panel analysis takes into account cross-sectional dependence. The long-run relationship between financial deepening and output emerges when financial and trade openness indices are included in the cointegrating relationship. The effects of different types of openness, however, are not uniform across developing and developed countries. Trade openness along with financial deepening are more important for the developing economies while financial deepening along with financial openness appear as more important for the advanced economies. Trade openness has the highest long-run coefficient and it is 33% higher for developing economies. No strong evidence of causality between financial development and output exists in the short-run. In the long-run there is clear evidence of causality from financial development to output. This causality is unidirectional in the advanced economies, corroborating the initial

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Schumpeterian view for the advanced economies. In developing countries, however, long run causality is bidirectional.

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Table 1
Definitions of Variables

Name	Definition	Source
Output (y)	Log of real per capita GDP	WDI-World Bank
Financial Development (fd)	Log of domestic credit provided by banks and financial institutions to private sector as percentage of GDP	WDI-World Bank
Financial Openness (fo)	Log of the stock of total flows of foreign assets and liabilities as percentage of GDP.	WDI-World Bank
Trade Openness (to)	Log of exports and imports as percentage of GDP	Lane & Milesi-Ferretti (2007)

Table 2
Panel Unit Root and Cross-sectional dependence Tests

(1) Group	(2) Variable	(3) IPS[a]	(4) CD[b]	(5) PES[c]
Advanced	Output	-2.126**	20.110***	-2.064
	Financial Development	-0.565	2.690***	-1.524
	Financial Openess	-0.174	25.390***	-2.202
	Trade Openness	-2.076**	42.320***	-2.167
Emerging	Output	1.331	9.040***	-1.606
	Financial Development	-0.466	4.160***	-2.129
	Financial Openess	-0.131	10.730***	-2.288**
	Trade Openness	0.067	10.130***	-1.705

Notes: a: IPS stands for the Im-Pesaran-Shin (2003) panel unit root test. Normal critical values are used.
b: CD stands for the Cross-sectional Dependence test of Pesaran (2004) which tests the null of independence against the alternative of dependence. Normal critical values are used.
c: PES stands for the Pesaran (2007) panel unit root test. Critical values are taken from Pesaran's Table 2.
***,**,* show rejection of null at 1%, 5% and 10%, respectively.

<p style="text-align:center">Table 3</p>
<p style="text-align:center">Chang and Song (2009) Panel Unit Root Tests</p>

Group	Variable	Average Tests			Minimum Tests		
		S^C	S^H	S^A	S^C_{min}	S^H_{min}	S^A_{min}
Advanced	Output	-0.976	1.010	0.917	-1.015	-0.330	-2.080
	Financial Development	1.937	-0.542	1.402	-1.175	-0.761	-1.107
	Financial Integration	1.554	1.225	-1.560	-0.871	-0.710	-2.205
	Trade Openness	1.366	0.268	0.035	-2.033	-1.834	-0.658
Emerging	Output	-0.902	0.595	0.023	-1.685	-1.452	-0.860
	Financial Development	-0.392	-0.616	-0.664	-1.931	-2.041	-1.986
	Financial Integration	2.194	-0.179	-0.544	-1.620	-1.083	-1.595
	Trade Openness	-0.912	0.748	0.329	-2.490	-1.792	-1.837

Note: Average tests are normally distributed. For minimum tests, critical values are taken from Chang and Song (2009).

<p style="text-align:center">Table 4</p>
<p style="text-align:center">Panel Cointegration Westerlund-Edgerton (2007) Bootstrap Test</p>

	Equation (1): $y_{i,t} = c_i + a_{1i} fd_{i,t} + u_{i,t}$	
	Asymptotic	Bootstrap
ALL	0.000	0.000
Advanced	0.007	0.000
Emerging	0.000	0.000

Note: Table reports asymptotic and bootstrap *p*-values of Westerlund-Edgerton test. For the bootstrap 1,000 replications were used.

<p style="text-align:center">Table 5</p>
<p style="text-align:center">Panel Cointegration Westerlund-Edgerton Bootstrap Test</p>

	Equation (2): $y_{i,t} = c_i + a_{1i} fd_{i,t} + \beta_{1i} fo_{i,t} + \gamma_{1i} to_{i,t} + v_{i,t}$	
	Asymptotic	Bootstrap
ALL	0.000	0.609
Advanced	0.000	0.862
Emerging	0.000	0.543

Note: Table reports asymptotic and bootstrap *p*-values of Westerlund-Edgerton test. For the bootstrap 1,000 replications were used.

Table 6

Panel Cointegrating Estimations

	Equation (2): $y_{i,t} = c_i + a_{1i}fd_{i,t} + \beta_{1i}fo_{i,t} + \gamma_{1i}to_{i,t} + v_{i,t}$		
	FD	FO	TO
ALL	0.219***	0.167***	0.373***
Advanced	0.156***	0.205***	0.324***
Emerging	0.294***	0.123***	0.431***

Note: Table reports the estimated coefficients for four equations using FMOLS. For the estimation Bartlett kernel is used. ***,**,* show statistical significance at 1%, 5% and 10%, respectively.

Table 7

Panel Causality Tests

	FD → GDP			GDP → FD		
	PMG-L[a]	PMG-S[b]	GM[c]	PMG-L[a]	PMG-S[b]	GM[c]
All	0.016	0.129	0.000	0.442	0.053	0.460
Advanced	0.000	0.351	0.000	0.385	0.074	0.390
Emerging	0.000	0.735	0.000	0.005	0.225	0.458

Note: Table reports the estimated *p*-values for panel causality tests.[a] PMG-L stands for the testing long-run causality using the Pooled Mean Group estimator. [b]PMG-S stands for testing the short-run causality. [c]GM presents the Group Mean test proposed by Canning and Pedroni (2008). In all cases, H_0 corresponds to the case of no causality, while H_1 indicates the existence of causality.