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In search of FDI-led growth in developing countries

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Abstract

This paper challenges the widespread belief that FDI generally has a positive impact on economic growth in developing countries. It addresses the limitations of the existing literature and re-examines the FDI-led growth hypothesis for 28 developing countries using cointegration techniques on a country-by-country basis. The paper finds that in the vast majority of countries FDI has no statistically significant long-run impact on growth. In very few cases, FDI indeed contributes to economic growth both in the long and the short run. But for some countries, there is also evidence of growth-limiting effects of FDI in the short or long term. Furthermore, our results indicate that there is no clear association between the growth impact of FDI and the level of per capita income, the level of education, the degree of openness, and the level of financial market development in developing countries.

JEL-Classification: F43; C22

Keywords: FDI; Growth; Developing countries; Cointegration

1. Introduction

Foreign direct investment (FDI) has grown dramatically in the past twenty years, exceeding the growth of the world production and the growth of international trade. Although most FDI is concentrated in the developed world, FDI flows have become increasingly significant for many developing countries. Since 1980, FDI to developing economies has increased over 12-fold (World Bank, 1999). Today, FDI typically accounts for more than 60 percent of private capital flows to the developing world (Carkovic and Levine, 2005; World Bank, 2006). This world-wide explosion of FDI was accompanied by a shift in emphasis among policymakers in developing countries to attract more foreign capital. Most countries have reduced barriers to FDI and many aggressively offered tax incentives and subsidies. The simple rationale for the increased efforts to attract FDI stems from believing that FDI promotes growth.

In theory there are several potential ways in which FDI can cause growth. For example, Solow-type standard neoclassical growth models suggest that FDI increases the capital stock and thus growth in the host economy by financing capital formation (Brems, 1970). Admittedly, in neoclassical growth models with diminishing returns to capital, FDI has only a "short-run" growth effect as countries move towards a new steady state (although the time frame involved in this adjustment can be quite long). Accordingly, the impact of FDI on growth is identical to that of domestic investment. In endogenous growth models, in contrast, FDI is often assumed to be more productive than domestic investment. The logic behind this is that FDI encourages the incorporation of new technologies in the production function of the host economy (Borensztein et al., 1998). In this view, FDI-related technological spillovers offset the effects of diminishing returns to capital and keep the economy on a long-term growth path. Moreover, endogenous growth models imply that FDI can promote long-run growth by augmenting the existing stock of knowledge in the host economy through labour training and skill acquisition, on the one hand, and through the introduction of alternative management practices and organisational arrangements, on the other (see, e.g., de Mello, 1997). In this context it is also argued that multinational companies, through FDI, may also diffuse their knowledge of global markets to domestic firms and hence enable them to become more successful exporters. In short, FDI is assumed to be an important vehicle for the transfer of technological and business know-how. These knowledge transfers may have substantial spillover effects for the entire economy. Hence, through capital accumulation and knowledge spillovers, FDI may play an important role for economic growth.

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Although the positive impact of foreign direct investment on economic growth seems to have recently acquired the status of a stylised fact (Campos and Kinoshita, 2002), a careful reading of the literature suggests that this positive relationship is far less definitive than generally believed. Agosin and Mayer (2000), for example, argue that FDI in the form of mergers and acquisitions do not necessarily increase the capital stock in capital-scarce economies. Cross-border mergers and acquisitions merely represent a transfer of existing assets from domestic to foreign hands. If the proceeds of the sales of these assets are spent on consumption, FDI does not contribute to capital formation and growth. This might be of particular relevance for many Latin American countries where a significant share of FDI flows in the 1990s occurred as a result of privatization of state-owned enterprises.

More importantly, the positive effect of FDI on growth through capital accumulation requires that FDI does not "crowd out" equal amounts of investment from domestic sources. Accordingly, FDI may actually harm the host economy when foreign investors claim scarce resources (such as import licenses, skilled manpower, credit facilities, etc.) or foreclose investment opportunities for local investors.

Additionally, there is also concern that the positive knowledge spillovers predicted by endogenous growth models do not occur in developing countries. For example, Görg and Greenaway (2004) critically review a number of firm-level studies on productivity spillovers in manufacturing industries in developing, developed and transition economies. They report that only six out of 25 studies using appropriate data and estimation techniques find some positive evidence of spillovers running from foreign-owned to domestic owned firms, none of which is for developing countries. One of these 25 ─ namely those by Aitken and Harrison (1999) for Venezuela ─ actually find some evidence of negative effects of the presence of multinationals firms.

Several explanations have been offered to explain these negative or statistically insignificant results. The most plausible explanation for the negative effects is that foreign firms reduce the productivity of domestic firms through competition effects, as suggested by Aitken and Harrison (1999). They argue that multinationals have lower marginal costs due to some firm specific advantage, which allows them to attract demand away from domestic firms, forcing them to reduce their production and move up their average cost curve. Furthermore, FDI is often associated with firm restructuring according to the production chain of the multinational company, which implies that raw materials and other inputs are purchased within the multinational enterprise and thus from foreign rather than local suppliers. As a consequence, the production of local suppliers may shrink.

There are also possible explanations for a failure to find any evidence for positive spillovers. Görg and Greenway (2004), for example, argue that multinationals may be able to effectively protect their firm-specific knowledge and, therefore, no knowledge spillovers occur. Moreover, domestic firms using very backward production technology and low skilled workers may be unable to learn from multinationals. And finally, knowledge spillovers are realised only if local firms have the ability to invest in absorbing foreign technologies. However, this ability may be restricted by underdeveloped local financial markets.[1]

Despite these potential negative effects, the empirical evidence generally suggests that FDI has a positive impact on economic growth in developing countries, as recent surveys by Lim (2001) and Hansen and Rand (2006) attest. Admittedly, the size of the impact of FDI on growth seems to depend on economic and political conditions in the host country, such as the level of per capita income, the human capital base, the degree of openness in the economy, and the extend of the development of domestic financial markets.

In this paper, we call this finding into question by, first, pointing to problems in the current empirical literature and, second, undertaking our own empirical assessment where we find little support for the growth-enhancing effect of FDI. Turning to the first issue, standard cross-country

[1] In addition to concerns that positive spillovers and capital accumulation effects do not occur in developing countries, some authors have raised concerns about possible balance of payments problems due to the repatriation of profits by foreign investors (Seabra and Flach, 2005) and harmful environmental impacts, if multinationals use FDI to "export" production no longer approved in their home countries (Wheeler, 2001).

and panel data studies on this topic implicitly assume that the relationship between FDI and growth is identical across countries. Significant differences between countries in the FDI-growth relationship can therefore lead to highly misleading results. Second, cross-country analyses often fail to address problems of reverse causality. Third, many cross-country studies use time-averaged data which, as shown below, can induce a spurious correlation. Forth, cross-country and panel data regression often restrict the relationship between FDI and GDP to be in growth rates or in first differences. This, however, precludes the possibility of a long-run or cointegrating relationship between the levels of the variables a priori. Fifth, indeed, time series analyses for individual countries generally investigate the FDI-growth nexus in the context of cointegration modelling that allows for a relationship between FDI and GDP both in levels and in first differences. But the vast majority of these studies uses only one approach to test for cointegration – namely the system-based cointegration procedure developed by Johansen (1995). This approach, however, has been shown to lead to severe biases in small samples (see, e.g., Reinsel and Ahn, 1988; Cheung and Lai, 1993).

This paper addresses each of these issues and then re-examines the impact of FDI on growth in developing countries. It contributes to the existing literature in the following ways: First, we analyze the data, country by country, to control for heterogeneity in the FDI-growth relationship between different countries. More concretely, we use time series techniques to investigate the relationship between FDI and growth for 28 developing countries individually. Second, we use the concept of Granger causality to investigate whether FDI actually leads to economic growth and not vive versa. Third, we use annual data rather than long-time averages to avoid the estimator bias due to temporal aggregation. Fourth, we allow for both long-run relationships between FDI and GDP in levels and short-run relationships between FDI and GDP in first differences. This approach entails a cointegration analysis of FDI and GDP, which also permits distinguishing between long-run and short-run Granger causality. Fifth, as far as the empirical methodology is concerned, we use the single equation cointegration tests proposed by Engle and Granger (1987) and Ericsson and

MacKinnon (2002) as well as the system technique advocated by Johansen (1995). The choice of using both single equation and system estimation methods is motivated by the work of Haug (1996) who finds that single-equation tests generally have smaller size distortions, but also lower power than system-based tests. He therefore recommends the application of both sets of tests. Because, however, standard cointegration procedures are biased towards accepting the null of no cointegration in the presence of structural breaks, we additionally use the Gregory and Hansen (1996) cointegration approach to allow for a potential structural break in the cointegration relation. Conditional on the presence of cointegration, we estimate the coefficients of the long-run relationships using Phillips and Loretan's (1991) nonlinear least squares method. This procedure has been shown to have good finite sample properties. Moreover, it generates unbiased and asymptotically efficient estimates for variables that cointegrate, even with endogenous regressors (Stock and Watson, 1993). Finally, for those countries whose FDI and GDP cointegrate, we investigate the direction of both long-run and short-run Granger causality by using a weak exogeneity testing procedure similar to the one proposed, among others, by Herzer et al. (2006). For those countries where no long-run relationships exist, we examine the direction of short-run causality by using standard Granger causality tests based on first differenced VARs.

Our main finding is that, in the vast majority of countries in our sample, FDI does not have a statistically significant long-run impact on GDP. Only in few cases, FDI contributes to long and / or short-run economic growth. In one country (Ecuador), we actually find evidence of a significant negative long-term effect of FDI on growth. Furthermore, our results indicate that there is no clear association between the growth impact of FDI and the level of per capita income, the level of education, the degree of openness, and the level of financial market development.

The reminder of the paper proceeds as follows. Section 2 reviews the empirical literature on this topic. Section 3 presents the empirical analysis. Concluding remarks are contained in Section 4.

2. The Empirical Literature: Review and Critique

A large body of empirical work examines the impact of FDI on economic growth in developing countries. This section reviews the main contributions and critiques the empirical methods employed therein. The discussion is organised around the econometric methodologies, which include cross-country regressions, dynamic panel and panel cointegration techniques as well as individual time series analyses.

2.1. Cross-Country Studies

Cross-country studies generally suggest a positive role for FDI in generating economic growth. However, the impact seems to be conditional on a number of factors, such as the level of per capita income, human capital, trade openness, and financial market development, as several studies show. For example, Blomström et al. (1994), using cross-country data from 78 developing countries, with averages taken over the period 1960-1985, find that lower income developing countries do not enjoy substantial growth benefits from FDI, whereas higher income developing countries do. The authors conclude from this that a certain threshold level of development is necessary to absorb new technology from investment of foreign firms.

Balasubramanyam et al. (1996), on the other hand ─ using cross-country data averaged over the period 1970-1985 for a sample of 46 developing countries ─ find that trade openness is crucial for acquiring the potential growth impact of FDI. They argue that more open economies are likely to both attract a higher volume of FDI and promote more efficient utilisation thereof than closed economies. Moreover, their estimates indicate that FDI has stronger effects on growth than domestic investment, which may be viewed as a confirmation of the hypothesis that FDI acts as a vehicle of international technology transfer.

Borensztein et al. (1998) use cross-country analysis of 69 developing countries with panel data over two separate time-periods 1970-1979 and 1980-1989. They find that the effect of FDI on growth depends on the level of human capital in the host country and that FDI has positive growth effects only if the level of education is higher than a given threshold.

And finally, Alfaro et al. (2004) examine the links among FDI, financial markets, and economic growth using cross-country data from 71 developing and developed countries averaged over the period 1975-1979. Their empirical evidence suggests that FDI plays an important role in contributing to economic growth, but the level of development of local financial markets is crucial for these positive effects to be realised. The logic behind this is that local firms generally need to reorganise their structure (buy new machines, and hire new managers and skilled labour) to take advantage of FDI-induced knowledge spillovers, which, however, is difficult if not impossible in underdeveloped financial markets.

Given the problems inherent to cross-country studies, these findings on FDI and growth should however be viewed sceptically. A major problem with cross-country studies is that they implicitly assume a common economic structure as well as similar production technologies and thus parameter homogeneity across countries. In fact, production technologies, institutions, and policies differ substantially, so that the effects of FDI on growth may also differ significantly between countries. The consequence of assuming parameter homogeneity is that cross-country regression results are not robust to the selection of countries (Ericsson et al., 2001). Moreover, country-specific effects due to omitted variables become part of the error term, which may result in biased estimates (Carkovic and Levine, 2005).

Apart from this, a statistically significant coefficient of FDI in the growth equation need not necessarily be the result of an impact of FDI on economic growth. Given that rapid economic growth usually generates higher demand and better profit opportunities for FDI, a positive correlation can also be compatible with causality running from growth to FDI (Nair-Reichert and Weinholt, 2001).

Additional problems arise when the data are averaged over time. Ericsson et al. (2001) demonstrate that temporal aggregation typically introduces simultaneity if causality in the sense of Granger (1996) exists. As a consequence, averaging data can induce a contemporaneous correlation between the time-averaged variables even if the original unaveraged variables are contemporaneously uncorrelated. Conversely, temporal aggregation can induce an apparent lack of relationship even if one exists.

2.2. Panel Studies

A solution to at least some of the problems discussed above is the use of panel data techniques. Panel estimation makes it possible to account for unobserved time-invariant country-specific-specific effects, therefore eliminating a possible source of omitted-variable bias. Furthermore, by including lagged explanatory variables, panel procedures allow to control for endogeneity bias. In the framework of dynamic model specifications it is also possible to explicitly test for Granger-causality. For example, Carkovic and Levine (2005) use the GMM dynamic panel data estimator of Arellano and Bover (1995) and Blundel and Bond (1998) with data averaged over seven 5-year periods between 1960 and 1995 for a sample of 68 countries. After controlling for the potential biases induced by endogeneity and omitted variables, they do not confirm the results of the cross-country studies by Blomström et al. (1994), Balasubramanyam et al. (1996), Borensztein et al. (1998), and Alfaro et al. (2004). Using econometric specifications that allow FDI to influence growth differently depending on national income, trade openness, education, and domestic financial development, Carkovic and Levine (2005) find that FDI does not exert a robust, positive impact on economic growth.

A serious problem with traditional panel data estimators, such as the one used by Carkovic and Levine (2005), however is the imposition of homogeneity on the coefficients of lagged dependent variables, when in fact the dynamics are heterogeneous across the panel. As Weinhold

(1999) and Nair-Reichert and Weinhold (2001) point out, this misspecification can lead to serious biases. To remedy this problem, Nair-Reichert and Weinhold (2001) use what they refer to as the mixed fixed and random (MFR) coefficient approach to test for causality between FDI and growth. The MFR allows for completely heterogeneity in the coefficients on the explanatory variables, thus avoiding the biases induced by imposing unrealistic homogeneity conditions on coefficients of the lagged dependent variables. Using data from 1971 to 1995 for 24 developing countries, they find that FDI on average has a significant positive impact on growth. But this relationship is very heterogeneous across countries.

It is problematic, however, that the standard cross-country and panel regressions on FDI and growth typically restrict the relationship between FDI and GDP to be in growth rates or first differences, since this precludes the possibility of a long-run or cointegrating relationship between the levels of the variables *a priori*. Furthermore, the conclusion, based on regressions in growth rates or first differences, that there is a statistically significant positive long-run association between FDI and GDP does not necessarily hold in the context of cointegration modelling that allows for a relationship between FDI and GDP both in levels and in first differences. Hence, as Ericsson et al. (2001) demonstrate, simply using first differences or growth rates can lead to serious misspecification problems ─ even in cross-cross country analyses.

2.3. Panel Cointegration Studies

In response to these criticisms, more recent econometric studies use panel cointegration techniques. For example, Basu et al. (2003) apply cointegration and causality tests to examine the issue of two-way causality between these two variables using a panel of 23 developing countries over the period 1978-1996. Allowing for individual country and time fixed effects as well as country-specific cointegration vectors they find a cointegrating relationship between FDI and GDP. Their results indicate bidirectional causality between these two variables for relatively open economies. For relatively closed economies long-run causality mainly runs from growth to FDI, implying that growth and FDI are not reinforcing under restrictive trade regimes.

Similarly, Hansen and Rand (2006) analyse the Granger-causal relationships between FDI and GDP in a sample of 31 developing countries for the period 1970-2000. Using estimators for heterogeneous panel data they find cointegration between FDI and GDP as well as between the share of FDI in gross capital formation and GDP. Their empirical evidence indicates that FDI has a lasting impact on GDP, whereas GDP has no long-run impact on FDI. They also find that a higher ratio of FDI in gross capital formation has positive effects on GDP. Hansen and Rand (2006) interpret this finding as evidence in favour of the hypotheses that FDI has an impact on GDP via knowledge transfers and adoption of new technologies.

However, with panel cointegration analyses, heterogeneity remains a serious concern. One problem is that a rejection of the null hypothesis of no panel cointegration can be driven by a few cointegrated relationships, as the findings by Gutierrez (2003) suggest. In addition, and perhaps more significantly, the presence of cointegration *across* countries in a panel may seriously bias upwards the probability of type I error of panel cointegration tests. That is, the null of no panel cointegration is falsely rejected too often, as Banerjee et al., (2004; 2005) show.[2] Consequently, there is a high risk that the whole panel may be erroneously modelled as cointegrated while only a few or even none of the (within-country) relationships are actually cointegrated.[3] Therefore, Banerjee at al. (2004) argue that many of the conclusions in the panel literature may be based upon misleading inference, and that it may be better to look at the evidence from country-by-country analyses.

[2] Existing panel cointegration tests generally do not account for the possibility of long-run cross-country dependence. They impose the restriction that there are no cointegrating relationships among the variables across the countries in the panel. Banerjee et al. (2004; 2005) show that this restriction is very likely to be violated in many macroeconomic time series because of economic links across countries. Therefore, any "automatic" use of panel methods for testing for cointegration might lead to wrong inferences.

2.4. Time Series Studies for Individual Countries

Several studies use cointegration techniques to investigate the causal relationship between FDI and growth for individual countries. For example, Zhang (2000) examines cointegration and causality between FDI and growth for 11 developing countries in East Asia and Latin America covering the period 1970-1995. His tests indicate cointegration and long-run Granger-causality from FDI to GDP for five countries. One out of the six countries without cointegration exhibits short-run Granger causality from FDI to growth. Cuadros et al. (2004), using quarterly data from 1980 to 2000, find cointegration between FDI and GDP for two out of three Latin American countries, where in these two countries long-run and short-run causality runs from FDI to GDP. Similarly, the results by Ramírez (2000) indicate for the period 1960-1995 that FDI Granger-causes GDP in Mexico, both in the short and in the long run. Xiaohui et al. (2002) use quarterly data for China from 1981 to 1997 and find cointegration as well as bi-directional short-run and long-run causality between FDI and GDP.

However, these results should also be viewed sceptically. Most of the time series studies on FDI and growth use only one method to test for cointegration – namely the system-based cointegration procedure developed by Johansen (1995), which may tend to falsely reject the null of no cointegration in small samples (see, e.g., Reinsel and Ahn 1988, Cheung and Lai, 1993). Therefore, it is not certain from the studies above that cointegration as well as causality between FDI and GDP is really supported by the data.

Consequently, the empirical literature on FDI and growth suffers from several shortcomings and must therefore be viewed with caution. Nevertheless, the available evidence, with the exception of Carkovic and Levine (2005), seems to suggest that FDI generally has a positive impact on economic growth in developing countries. This conclusion is confirmed by several surveys on this topic, as already noted in the beginning of the paper. Admittedly, the impact seems to be country

[3] The mixing of cointegrated and non-cointegrated relationships can lead to serious biases in determining causality as well as the long-run parameters, if only a small fraction of the relationships in the panel actually cointegrates (see, e.g.,

specific, depending on factors, such as the level of per capita income in the host country, the human capital base, the degree of openness in the economy, and the level of financial market development. In the next section, we re-examine these issues using time series for individual countries.

3. Empirical Analysis

In this section, we empirically investigate the relationship between FDI and growth for 28 developing countries using several single-equation and system cointegration techniques. In particular, we are concerned with the following questions:

1. In how many countries are there cointegrating or long-run relationships between FDI and GDP and what are the parameters or coefficients of these relations?
2. What is the direction of causality? Does FDI lead to long-run growth or vice versa?
3. Is there short-run Granger causality between FDI and GDP in those countries whose FDI and GDP are not related in the long run?
4. Is there a link between the impact of FDI on GDP in developing countries and their level of per capita income, human capital, trade openness, and financial market development?

3.1. Variables and Data

To answer these questions we rely on the standard time series and panel cointegration literature on FDI and growth and use a bivariate model (see, e.g., Zhang, 2001; Basu et al., 2003; Hansen and Rand, 2006; Chowdhury and Mavrotas, 2006). Following Hansen and Rand (2006), the variables considered for the model are the following: $Log\, GDP_t$ and FDI_t / GDP_t, where the former denotes the logarithm of real GDP (in constant 2000 US dollars)[4] at time *t*, and the latter is the FDI-to-GDP ratio (in percent) at time *t*. Note that we do not use (log) FDI but the FDI-to-GDP ratio. The

Strauss and Wohar, 2004).

[4] We follow the standard in the literature (see, e.g., Hanson and Rand, 2006; Alfaro et al. 2004. Nair-Reicher and Weinhold, 2001; Basu et al. 2003) and use real GDP in constant US$ using official exchange rates as contained in the

reason is that FDI, via the national income accounting identity, is itself a component of GDP and thus partly endogenous within the GDP equation, which may bias the results in favour of a correlation between these two variables. Therefore, as in many previous studies (see, e.g., Nair-Reichert and Weinhold, 2001; Hansen and Rand, 2006), we use the FDI-to-GDP ratio.

The data utilised in this study are annual data from the World Development Indicators CD-ROM (WDI 2005). Following previous research (see, e.g., Nair-Reichert and Weinhold, 2001; Basu et al, 2003; Hansen and Rand, 2006), we employ net FDI, defined as net inflows of investment to acquire a lasting management interest (10 per cent or more of voting stock) in an enterprise operating in an economy other than that of the investor. It includes equity capital, reinvestment of earnings, and other long term and short-term capital as shown in the balance of payments.

Our sample consists of 28 developing countries. The sample is very similar to that of Nair-Reichert and Weinhold (2001), who use data for a sample of 24 developing countries, Basu et al. (2003), who employ data for 23 developing countries, and Hansen and Rand (2006), who use data for 31 developing countries. Out of the 28 countries we investigate, 10 are in Latin America, i.e. Argentina, Brazil, Costa Rica, Ecuador, Dominican Republic, Mexico, Chile, Peru, Venezuela, Columbia, 9 are in Asia, i.e. Sri Lanka, India, Pakistan, Indonesia, Malaysia, Korea, Philippines, Thailand, Singapore, and 9 are in Africa, i.e. Cameroon, Zambia, Côte d'Ivoire, Nigeria, Tunisia, Kenya, Morocco, Ghana, Egypt. According to the World Development Indicators (2005), these countries are major recipients of FDI in their respective regions. Hence, we believe that the countries selected provide a fair representation of the major developing areas in the world. A more complete picture of the developing world would, of course, require the inclusion of many other countries, in particular China, which is the largest recipient of FDI flows among developing countries. But the availability of data for a reasonably long time period (at least 25 years) limits our

World Development Indicators. We also redid the analysis using real growth using local currencies and found qualitatively the same results.

choice to these 28. For all but two countries ─ Korea and Singapore ─ the sample period is 1970-2003. The sample period for Korea is 1976-2003 and 1972-2003 for Singapore.

Table 1 gives an impression of the evolution of the variables in the sample periods. From Table 1, it can be inferred that all series exhibit non-stationary behaviour. The GDP (in logarithms) grew modest to strong in all countries involved. The largest growth dynamic is observed in Singapore, the lowest in Zambia. Similarly, in almost all countries the FDI-to-GDP ratio increased significantly between the early 1970s and the late 1990s. Exceptions are Kenya and Indonesia, with a declining trend. In Cameroon the FDI-to-GDP ratio indeed increased, but only moderately. Moreover, from Table 1, it can be seen that the importance of FDI differs dramatically between countries. In Singapore, for example, average net FDI inflows account for 15.05 percent of GDP in the period 1999-2003. In contrast, for Indonesia the share of net FDI flows in GDP (-1.69%) is actually negative due to the repatriation of foreign capital.

[Table 1 around here]

Not evident in Table 1, but also noteworthy is that the FDI-to-GDP ratio is highly volatile compared to the (log) GDP, in particular in African countries. Furthermore, several FDI-to-GDP series contain structural discontinuities. Therefore, in order to determine the time series properties of the variables we have carried out standard ADF tests as well as Perron unit root tests allowing for structural breaks in the series. The results of these tests (not reported here to save space) indicate that all series are non-stationary in levels, but stationary in their first differences. In the following, we therefore treat $Log\,GDP_t$ and $FDI_t\,/\,GDP_t$ as integrated of order one, $I(1)$. Thus, the next step in our analysis is an investigation of the cointegration properties of these series.

3.2. Testing for Cointegration: The Engle-Granger, the ECM, and the Johansen Approach

There are several methods to test for cointegration. The most commonly used are (*i*) the single-equation Engle-Granger (1987) two-step procedure, (*ii*) the single-equation conditional error correction model (ECM) test initially proposed by Phillips (1954) and further developed, among others, by Banerjee et al. (1998), Pesaran et al. (2001), and Ericsson and MacKinnon (2002), and (*iii*) the system based cointegration approach of Johansen (1995). As noted in the beginning of this paper, single-equation tests tend to have smaller size distortions, but also lower power than system-based tests (Haug, 1996). Of course, each method has specific advantages and disadvantages, since it imposes particular assumptions about the data generating process. Hence, in order to draw robust conclusions about the long-run relationship between FDI and growth, we use all three approaches.

(*i*) *The Engle-Granger Procedure*

Following the Engle-Granger procedure, we estimate the static cointegrating regression

$$Log\,GDP_t = c_1 + \delta_1\, t + \beta_1\,(FDI_t\,/\,GDP_t) + e_t\,, \qquad t = 1, 2, ..., T\,, \tag{1}$$

where β_1 is the long-run coefficient, c_1 is a constant parameter, and e_t is the usual error term. At least initially, a linear trend t with coefficient δ_1 is included in equation (1), because most of the data show steadily rising trends. In cases where the inclusion of a linear time trend is not supported by the data, we use specifications with only a constant term.

In the second step, we test the estimated residuals, $\hat{e}_t$, for stationarity by estimating the ADF test regression

$$\Delta\hat{e}_t = \rho\hat{e}_{t-1} + \sum_{j=1}^{k}\beta_j\Delta\hat{e}_{t-j} + v_t\,, \tag{2}$$

where k is the lag length. Following Herzer et al. (2006) the lag length in the ADF regressions is determined using the t-sig method, i.e. downward testing beginning with an arbitrarily large number of lags ─ in our analysis five. If the last included lag is insignificant, the number of lags is reduced by one and the test regression is re-estimated until a significant lagged variable is found.

The null hypothesis of no cointegration is rejected, if the t-statistic of the ρ coefficient is in absolute value greater than the finite sample critical values obtained from the response surfaces of

MacKinnon (1991).[5] Important to note is that weak exogeneity of the regressors is indeed assumed but not required for the Engle-Granger approach, as pointed out by Ericsson and MacKinnon (2002). However, the Engle-Granger approach imposes a common factor restriction on the dynamics of the relationship between the variables, which is only appropriate when the short-run coefficients equal the long-run coefficients (see, e.g., Kremers et al., 1992). Consequently, if the short-run and the long-run coefficients markedly differ, a loss of power relative to the ECM and Johansen procedures may result. Therefore, we use the ECM and the Johansen cointegration procedure in addition to the Engle-Granger approach.

(*ii*) *The ECM Procedure*

The ECM procedure is based on a conditional error correction model, which in our case is given by

$$\begin{aligned} \Delta LGDP_t = {} & b_1 + b_2 t + b_3 LGDP_{t-1} + b_4 (FDI_{t-1} / GDP_{t-1}) \\ & + \sum_{i=1}^{k} \eta_i \Delta LGDP_{t-i} + \sum_{i=0}^{k} \gamma_i \Delta (FDI_{t-i} / GDP_{t-i}) + u_t . \end{aligned} \tag{3}$$

Following Wolters et al. (1998), the lag length k is determined using the general-to-specific approach. Concretely, we allow for up to two lags before successively eliminating the variables with the lowest t-values from the short-run dynamics of the equation. Also, the time trend t is excluded from the equation, if it turns out to be insignificant. A significant negative coefficient of the lagged dependent level variable indicates cointegration. Accordingly, the null of no cointegration to be tested is $b_3 = 0$. To this end, Ericsson and MacKinnon (2002) provide an extensive set of critical values with finite sample corrections based on response surfaces similar to those in MacKinnon (1991) for the Engle-Granger procedure. Note that equation (3) does not impose the potentially invalid common factor restriction implied by the Engle-Granger framework. However, the ECM procedure assumes weak exogeneity of the explanatory variables and long-run endogeneity of the dependent variable, respectively, which may be empirically incorrect. This

[5] The response surfaces in MacKinnon (1991) allow construction of critical values with finite sample adjustments.

criticism does not apply to the full information maximum likelihood cointegration approach developed by Johansen (1995).

(*iii*) *The Johansen Procedure*

The Johansen procedure is based on a vector error correction model given by

$$\Delta y_t = \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \alpha\beta' y_{t-1} + \psi D_t + \varepsilon_t \,, \tag{4}$$

where y_t is a $n \times 1$ vector of endogenous variables ($y_t = [\,Log\,GDP_t,\ FDI_t\,/\,GDP_t]'$), β is a $n \times r$ matrix whose r columns represent the cointegrating vectors among the variables in y_t, α is a $n \times r$ matrix whose n rows represent the error correction coefficients, Γ_i is a $n \times r$ matrix of short-run coefficients, and ψ represents a $n \times r$ matrix of coefficients on D_t – a vector of deterministic terms, such as a constant term and a trend. In order to test for cointegration, we use the trace test, which tests the rank r of the $n \times n$ product matrix $\alpha\beta'$ such that the reduced rank, $r < n$, implies cointegration.[6] Osterwald-Lenum (1992) provides the critical values for this test.

However, as already noted, the Johansen procedure may tend to falsely reject the null of no cointegration in small samples. To alleviate this problem, we adjust the test statistics downward by using the small sample correction factor (T - $n \times k$) / T, as suggested by Reinsel and Ahn (1992). Another problem of the Johansen method is that the results are very sensitive to the specification of the statistical model and the choice of the lag length (Stock and Watson, 1993). In order to determine the appropriate lag length, we use the Schwarz criterion. The Schwarz criterion has been shown to lead to consistent estimates in both stationary and nonstationary models while the Akaike criterion is characterized by a positive limiting probability of overfitting (see, e.g., Pötscher, 1989, 1990). To find the most appropriate deterministic components, we rely on the Pantula Principle. This procedure involves moving from the most restrictive model (i.e., no deterministic components) to less restrictive assumptions (i.e., unrestricted intercept and restricted trend), comparing the test statistics for each model with the corresponding critical values, stopping only when the null of no cointegration is not rejected (see, e.g., Johansen, 1995).

(*iv*) *Results*

Table 2 reports the results of the Engle-Granger, the ECM, and the Johansen cointegration test. As can be seen, all three tests fail to reject the null of no cointegration in 24 of the 28 countries. Only for Ecuador and Sri Lanka, we can clearly reject the null hypothesis. For Mexico and Venezuela, the results are less clear. The Engle-Granger and the Johansen test reject the null of no cointegration whereas the ECM test does not. Admittedly, the failure of the ECM test to detect cointegration might be caused by the long-run endogeneity of the FDI to GDP ratio. To test this possibility, we re-run the ECM test with FDI_t / GDP_t as the dependent variable. For Mexico the estimated ECM t-statistic then amounts to -5.28, for Venezuela to -4.20. Comparing these values with the corresponding critical values in Table 2, the null hypothesis of no cointegration can be rejected at the 1% significance level. This finding suggests that FDI_t / GDP_t can be treated as endogenous. Consequently, long-run Granger-causality might be running from $Log\, GDP_t$ to FDI_t / GDP_t in Mexico and Venezuela. The issue of causality will be taken up later on.

[Table 2 around here]

At this point, we briefly return to the already mentioned problem of panel cointegration studies. As noted in Subsection 2.4, with panel cointegration procedures, there is a high risk that the whole panel may erroneously be modelled as cointegrated while only a small fraction of relationships are actually cointegrated. To demonstrate this, we also report the results of the heterogeneous panel cointegration test proposed by Larsson et al. (2001). This test is based on the above described Johansen (1995) procedure, where the panel test statistic, Ψ, is given by the

[6] Because equation (4) contains only two $I(1)$ Variables, n = 2, the cointegration rank in our case cannot exceed one. If

standardised average, LR, of the N (= 28) individual trace statistics: $\Psi = \left\{\sqrt{N}[LR - E(Z)]\right\} / \sqrt{Var(Z)}$, with $E(Z)$ and $Var(Z)$ as the mean and variance of the asymptotic trace statistic reported by Larsson et al. (2001). Larsson et al. (2001) show that this panel standardised cointegration trace statistic converges to a standard normal distribution N(0, 1). Accordingly, the 1% critical value is 2.326. Since Ψ is estimated to be 6.601, the null of no panel cointegration can be rejected at the 1% significance level.[7] Consequently, the panel cointegration test (falsely) indicates cointegration between $Log\, GDP_t$ and FDI_t / GDP_t for the whole panel. In contrast, the individual country-by-country test results in Table 2 clearly indicate that the vast majority of relationships are not cointegrated. As already mentioned, this large mismatch can be attributed (i) to the fact that the rejection of the null of no panel cointegration can be driven by a few cointegrated relationships (see, e.g., Gutierrez, 2003; Strauss and Wohar, 2004), and / or (ii) to a bias in the panel tests, that leads to the rejection of the null of no panel cointegration to often when there are cross-country cointegrating relationships (Banjerjee et al., 2004, 2005). Although this was not the focus of our study, we found indeed some instances for cross-country cointegration among some of the GDP series and some of the FDI-to-GDP series. (Results are available on request.)

However, one should still be cautious in accepting the null of no cointegration. The inability to reject the null of no cointegration in the large majority of countries might be due to the existence of structural breaks that bias the results. For a robustness check, we therefore allow for possible structural breaks by applying the Gregory-Hansen (1996) cointegration procedure to those countries for which we were not able to find cointegration using the Engle-Granger, the ECM, and the Johansen approach.

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the matrix $\alpha\beta'$ has the full rank $r = n$, this would imply that all n components of y_t are stationary.

3.3. Testing for Cointegration: The Gregory-Hansen Approach

We apply the single equation approach of Gregory and Hansen (1996) and not a system based approach, because structural breaks can be modelled more directly using single equations. Moreover, the Gregory-Hansen cointegration procedure allows for an unknown structural break whereas system based approaches usually require the breakpoint to be known *a priori*. Gregory and Hansen present the following models:

the level shift model (C)

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^T y_{2t} + e_t, \tag{5}$$

the slope change model (C/T)

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t, \tag{6}$$

and the regime shift model (C/S)

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t} \varphi_{t\tau} + e_t, \tag{7}$$

where in our analysis $y_{1t} = LogGDP_t$ and $y_{2t} = FDI_t / GDP_t$, μ_1 and α_1 are intercept and slope coefficients before the shift, μ_2 and α_2 denote changes to the intercept and slope coefficients at the time of the shift. The dummy variable $\varphi_{t\tau}$ is defined by

$$\varphi_{t\tau} = \begin{cases} 0, & if \quad t \le [\eta\tau] \\ 1, & if \quad t > [\eta\tau], \end{cases} \tag{8}$$

where the unknown parameter $\tau \in (0, 1)$ denotes the relative timing of the break, and $[\quad]$ denotes the integer part.

The breakpoint is determined by estimating the models for each possible break date in the data set, saving the estimated residuals from each iteration and selecting τ as the value which minimises the unit root test statistics for the estimated residuals. Accordingly, using the ADF

[7] For the panel cointegration test, the time span was determined by the shortest span for an individual variable in a particular country, i.e., 1976-2003. The individual trace statistics are adjusted for small sample bias using the Reinsel and Ahn (1992) method. Model assumption: unrestricted intercept and restricted trend.

statistics ─ calculated by estimating equation (2) ─ the statistics of interest are the smallest ADF t-values across all values of $\tau \in T$ and thus

$$ADF^* = \inf_{\tau \in T} ADF(\tau) \quad . \tag{9}$$

If ADF^* exceeds in absolute value the critical values reported by Gregory and Hansen (1996), the null of no cointegration is rejected. Table 3 presents the results.

[Table 3 around here]

As can be seen, the Gregory-Hansen test rejects the null hypothesis of no cointegration for Nigeria, Tunisia, and Egypt using the C/T model. In contrast, for Argentina, Brazil, Costa Rica, Dominican Republic, Chile, Peru, Columbia, India, Pakistan, Indonesia, Malaysia, Korea, Philippines, Thailand, Singapore, Cameroon, Zambia, Côte D' Ivoire, Kenya, Morocco, and Ghana we find no cointegration between FDI_t / GDP_t and $Log\, GDP_t$ even if we allow for possible structural breaks. Consequently, the cointegration analysis, even when allowing for structural breaks, indicates that there is no long-run relationship between FDI_t / GDP_t and $Log\, GDP_t$ in 21 out of 28 countries, and thus in 75 percent of cases.

3.4. Estimating the Long-Run Relationships

Having found that for Ecuador, Mexico, Venezuela, Sri Lanka, Nigeria, Tunisia, and Egypt FDI_t / GDP_t and $Log\, GDP_t$ are cointegrated, the next step in our analysis is the estimation of the long-run relationship between these variables. To this end, we use the Phillips and Loretan (1991) procedure. This procedure is asymptotically equivalent to Johansen's maximum likelihood estimator, but more robust to many particulars of the marginal process, such as specific lag lengths, especially in small samples (Stock and Watson, 1993). Moreover, by including leads and lags of the first differences of the explanatory variables the Phillips-Loretan approach generates unbiased and asymptotically efficient estimates, even in the presence of endogenous regressors. Additionally, the approach deals with the autocorrelation of the residuals by including lagged values of the stationary deviation from the cointegrating relationship. The Phillips-Loretan equation, which is estimated by nonlinear least squares, is given by

$$LogGDP_t = \lambda_1 ' y_{1_t} + b_1 (LogGDP_{t-1} - \lambda_1 ' y_{1_{t-1}}) + \sum_{i=-k}^{i=k} \Phi_{1_i} \Delta (FDI_{t+i} / GDP_{t+i}) + \varepsilon_t , \qquad (10)$$

where y_{1_t} denotes the vector $[1, t, FDI_t / GDP_t, \varphi_{t\tau}]'$, λ_1 denotes the corresponding coefficient vector $[c_1, \delta_1, \beta_1, \mu]'$, and Φ_{1_i} are coefficients of lead and lag differences of FDI_t / GDP_t.

We estimate Equation (10) for Ecuador, Sri Lanka, Nigeria, Tunisia, and Egypt, where a step dummy variable $\varphi_{t\tau}$ with coefficient μ is included for Nigeria, Tunisia, and Egypt to model the structural breaks detected by the Gregory-Hansen procedure. For Nigeria the dummy variable is one from 1982 onwards and otherwise zero. It presumably captures the effects of the dramatic fall in oil prices that plunged Nigeria into deep recession in 1982. For Tunisia the dummy takes the value one for the period after 1976 and zero before. Possible reasons for the importance of this dummy variable are the fifth development plan initiated by the Tunisian government in 1977, the free trade agreement of 1976 between Tunisia and the EU, and/or the massive increase in oil prices due to the oil crises of 1973-75, which led to rapid export earnings and GDP growth. For Egypt, the step dummy is one for the 1980-2002 period and zero otherwise. It controls for the economic growth that presumably was the result of long-term peace with Israel and sharply increasing aid inflows.

For Mexico and Venezuela we proceed under the assumption that FDI_t / GDP_t can be treated as endogenous, as suggested by the ECM tests in Section 3.3. Hence, the estimating equation for these two countries is given by

$$FDI_t / GDP_t = \lambda_2 ' y_{2_t} + b_2 (FDI_{t-1} / GDP_{t-1} - \lambda_2 ' y_{2_{t-1}}) + \sum_{j=-k}^{j=k} \Phi_{2_i} \Delta LogGDP_{t+j} + \varepsilon_t , \quad (11)$$

where y_{2t} and λ_2 denote the vector $[1, LogGDP_t]'$ and the vector $[c_2, \beta_2]'$, respectively. A linear trend was not supported by the data and hence was excluded from the equation.

The estimates of λ_1 and λ_2, along with their t-statistics as well as some residual diagnostics of equation (10) and (11) are reported in Table 4. The numbers in parentheses below the values of the diagnostic test statistics are the corresponding *p*-values. *LM*(*k*), *k* = 1, 3 are Lagrange Multiplier (LM) tests for autocorrelation based on 1 and 3 lags, *ARCH*(*k*) is an LM test for autoregressive conditional heteroscedasticity of order *k*, *k* =1 and 3, and *JB* is the Jarque-Bera test for normality. As can be seen, all *p*-values exceed the conventional significance levels. Hence, we conclude that the residuals do not show any signs of non-normality, autocorrelation or conditional heteroscedasticity.

[Table 4 around here]

Looking at the coefficients of FDI_t / GDP_t, it immediately becomes clear that the implied long-run impact of FDI on GDP varies strongly between the four countries. For Ecuador the long-run impact is actually negative. The Phillips-Loretan estimate of β_1 is -0.036, implying that a one percentage point increase in FDI_t / GDP_t leads to a 3.6 per cent *decrease* in GDP in the long run. In contrast, for Sri Lanka, Nigeria, Tunisia, and Egypt output *increases* by 9.8, 4.9, 2.7, and 4.7 percent, respectively, due to a one percentage point increase in FDI_t / GDP_t.

Treating FDI_t / GDP_t as endogenous, the estimated β_2 coefficient in Table 4 suggests that in Mexico a one percent increase in GDP is associated with a 0.06037 percentage point increase in FDI_t / GDP_t. From the estimated β_2 coefficient for Venezuela it can be inferred that the FDI-to-GDP ratio increases by 0.1093 percentage points in response to a one percent increase in GDP.

However, these interpretations are implicitly based on the assumption that long-run causality runs from $Log\,GDP_t$ to FDI_t / GDP_t in Mexico and Venezuela, and from FDI_t / GDP_t to

$Log\,GDP_t$ in Ecuador, Sri Lanka, Nigeria, Tunisia, and Egypt respectively. In order to investigate whether this assumption actually holds, we apply the concept of Granger causality.

3.5. Testing for Causality

Cointegration implies Granger-causality in at least one direction (see, e.g. Granger, 1988). For those countries with cointegration between FDI_t / GDP_t and $Log\,GDP_t$, we investigate the direction of both long and short-run causality by using a weak exogeneity testing procedure similar to the one proposed, among others, by Herzer et al. (2006). For the countries whose FDI and GDP variables are not related in the long run, the two variables can affect each other in the short run. To investigate this, we apply the standard Granger-causality test using VAR models in first differences.

(*i*) *Long-Run Causality*

The causality testing procedure for Ecuador, Sri Lanka, Nigeria, Tunisia, Egypt, Mexico, and Venezuela involves estimating a vector error correction model given by

$$\begin{bmatrix} \Delta LogGDP_t \\ \Delta(FDI_t / GDP_t) \end{bmatrix} = \begin{bmatrix} \upsilon_1 \\ \upsilon_2 \end{bmatrix} + \sum_{k=1}^{p-1} \Gamma_k \begin{bmatrix} \Delta LogGDP_{t-k} \\ \Delta(FDI_{t-k} / GDP_{t-k}) \end{bmatrix} + \begin{bmatrix} a_1 \\ a_2 \end{bmatrix} ec_{t-1} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}, \tag{12}$$

where the error correction term

$$ec_t = LogGDP_t - [\hat{c}_1 + \hat{\delta}_1 t + \hat{\beta}_1 (FDI_t / GDP_t) + \hat{\mu}\varphi_{t\tau}] \quad \text{and} \tag{13}$$

$$ec_t = FDI_t / GDP_t - [\hat{c}_2 + \hat{\beta}_2 LogGDP_t], \tag{14}$$

respectively, is the residual of the Phillips-Loretan long-relations in Table 4. According to Granger (1988) a significant error correction term indicates long-run Granger causality from the independent to the dependent variables, where long-run Granger non-causality and weak exogeneity can be regarded as equivalent (see, e.g., Hall and Milne, 1994). Note that here we are not interested in the short-run Granger causality, but in the long-run effects. Therefore, at this stage, we do not aim to test for the joint significance of the lagged differenced variables, i.e., the short-run dynamics. The issue of short-run causality is taken up below. Following Herzer et al. (2006) we test for weak exogeneity and thus for long-run Granger non-causality by eliminating the short-run parameters (Γ_k) successively according to the lowest t-values and then we decide on the significance of the error correction term. In doing so, we reduce the number of parameters and thereby we increase the precision of the weak exogeneity tests on the α coefficients. Because all variables in (12), including ec_{t-1}, are $I(0)$ variables, conventional t-tests can be used. Due to the small sample size, we allow for a maximum number of two lags. After applying the general-to-specific model reduction procedure, we obtain the results in Table 5.

[Table 5 around here]

According to the t-statistics in Table 5, the null hypothesis of weak exogeneity of both $Log\, GDP_t$ and FDI_t / GDP_t is rejected for Sri Lanka, Nigeria, Tunisia, and Egypt. Thus, the weak exogeneity tests suggest a long-run feedback relationship between FDI and GDP since both variables can be regarded as endogenous. In other words, Sri Lanka, Nigeria, Tunisia, and Egypt exhibit long-run causality between FDI and growth running in both directions. In contrast, we cannot reject weak exogeneity of $Log\, GDP_t$ for Mexico and Venezuela. Hence, we conclude that in Mexico and Venezuela FDI is a consequence of economic growth rather than the cause.[8]. Ecuador is the only out of 28 countries for which we find unidirectional long-run Granger causality from FDI_t / GDP_t to $Log\, GDP_t$. However, the output of Ecuador is negatively affected by FDI in the period under investigation, as indicated by the Philips-Loretan estimates in Table 4. Consequently, in the long run, growth is positively affected by FDI in only four out of 28 countries.

(*ii*) *Short-run Causality*

For 21 of the 28 countries we had found no cointegrating relationship between FDI_t / GDP_t and $Log\, GDP_t$. The absence of cointegration implies the absence of long-run Granger causality, but it does not preclude causality in the short run. To test for short-run Granger causality in the non-cointegrated countries, we apply the standard Granger-causality test using a VAR in first differences. To test for short-run Granger causality in the cointegrated countries, we again use a vector error correction model. In particular, for those countries without cointegration we estimate equation (12) without including the error correction term, for those countries with cointegration we estimate equation (12) with ec_{t-1}, and examine the joint significance of the short-run coefficients Γ_k by means of an *F*-Test. Following Miller (1991), the lag structure is determined using the general-to-specific approach. We estimate the VAR with up to two lags and eliminate lags with insignificant coefficients according to the lowest *t*-values. Table 6 reports the results.

[Table 6 around here]

The table suggests that there is no causality (neither in the short nor in the long run) between FDI and growth in Argentina, Brazil, Costa Rica, the Dominican Republic, Pakistan, Korea, the Philippines, Singapore, Cameroon, Côte d'Ivoire, and Morocco. In the short run, unidirectional causality runs from FDI_t / GDP_t to $Log\, GDP_t$ in Chile, Peru, Columbia, India, Zambia, Kenya, Sri Lanka, and Egypt. However, in the short run, output is negatively affected by FDI in Columbia, Zambia, Sri Lanka, and Egypt as indicated by the coefficients of $\sum \Delta(FDI_t / GDP_t)$ in the $\Delta LogGDP_t$ equation. Accordingly, in Sri Lanka as well as in Egypt FDI has negative short run but positive long run effects on GDP (see the Phillips Loretan estimates in Table 4). For Indonesia, Ghana, and Tunisia, we find unidirectional short-run causality running from $Log\, GDP_t$

[8] We speculate that this effect might be related to the fact that both countries are oil exporters, where growth is affected by oil exports and where FDI in the oil-extracting sectors surge when oil-induced growth is high.

to FDI_t / GDP_t. Only Thailand exhibits bidirectional short-run Granger causality between the FDI to GDP ratio and economic growth.

Summarising, we find that FDI has a long-run positive impact on GDP in 15 percent of the countries in our sample. For one out of 28 countries ─ and, thus, 3.6 percent of cases ─ we find long-run negative effects of FDI on GDP. In the short-run, FDI positively affects growth in about 18 percent of cases. In 15 percent of countries in our sample, the short-run impact of FDI on GDP is negative. Note that we checked the robustness of these results (i) by using several alternative methods for estimating the cointegrating parameters, and (ii) by including exports as a control variable in our models, as suggested, for example, by Cuadros et al. (2003). The results (not reported here, but available on request) do not change qualitatively.

Against this background, we cannot conclude ─ as many studies do ─ that FDI has a generally positive impact on economic growth in developing countries. In contrast, in the vast majority of countries in our sample, FDI has no statistically significant long-run impact on GDP. In few countries, FDI contributes to long and / or short-run economic growth. But for some countries, we also find strong evidence of growth-limiting effects of FDI in the long or short term. Furthermore, the estimation results in Table 4-6 seem to suggest that there are no systematic differences in the growth impact of FDI between Latin American, Asian, and African countries. Although this is not the prime concern of our study, it appears that the impact of FDI on growth is more or less randomly distributed across regions.

3.6. Searching for a Link between the Impact of FDI on GDP and Country-Specific Factors

As noted in section 2, several studies suggest that the impact of FDI on economic growth depends on country-specific factors, such as the level of per capita income in the host country, the human capital base, the degree of openness in the economy, and the level of financial market development. If this is really the case, then the country-specific impact of FDI on GDP should vary systematically with these factors. To investigate this, we consider four simple scatter plots of the estimated long-run impact of FDI on GDP (from Table 4) against the period-averages of log of GDP per capita, secondary enrolment rates, imports plus exports as a percentage of GDP, and domestic credit to the private sector as a percentage of GDP.[9] Note that the long-run impact of FDI on GDP was set to zero for those countries that do not exhibit long-run Granger causality from FDI to GDP.

[Figure 1 around here]

As can be seen from Figure 1, there is no clear (linear) relationship between the estimated long-run growth impact of FDI and either one of these indicators. Additionally, in Figure 2 we plot the short-run impact of FDI on GDP (from Table 5) for each country against the log GDP per capita, the trade volume, the secondary schooling rate, and domestic credit for each country. Again, no clear relationship is found.

[Figure 2 around here]

Of course, our sample is too small to draw statistically valid inferences. Yet, our results seem to support Carkovic and Levine (2005), who also found that the impact of FDI on growth does not robustly vary with the level of per capita income, human capital, trade openness, and financial market development. On the other hand, we differ from Carkovic and Levine (2005) in that we also find growth-limiting effects of FDI.

4. Concluding Remarks

In this paper we call into question the growth-inducing effect of FDI by first pointing to the weaknesses in empirical literature on this issue and then undertaking our own empirical analysis that generally finds little support for such an effect. In the vast majority of countries there is neither a long-term not a short-term effect and our results do not indicate a clear regional pattern or the influence of other factors on the FDI-growth linkage.

[9] All data are from the World Development Indicators (2005).

We speculate that the following reasons might account for this failure to find a relationship. First, FDI as a share of GDP is, particularly in the 1970s and 1980s, rather small, often amounting to less than 1% of GDP and thus also constituting only an insignificant share of total investment. Thus FDI might simply be too marginal to have a serious impact. Given the rising prominence of FDI in the 1990s, it might be worthwhile to periodically redo the time series analysis undertaken here to see whether the low levels of FDI are really the reason for the failure to find an effect. Given that most FDI usually goes to countries that already have substantial domestic savings rates (OECD, 2002), it remains unclear, however, whether even rising FDI inflows really play such a substantial role.

Second, it could well be that the growth-limiting effects of FDI often limit the growth-enhancing effects leading to little net impact. As discussed at the beginning of the paper, there are a range of possible factors that ensure that FDI promotes or hinders economic growth. The factors are likely to differ between countries and between types of FDI and sectors of destination. For example, the effects of FDI in manufacturing might differ from those in extractive sectors which again might differ from FDI that is a result of privatization of state-owned enterprises.

Thus we suggest that future research should focus on identifying the types of FDI that might promote growth in particular circumstances, rather than presuming that a positive effect generally exists.

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Table 1

GDP (in logarithms) and FDI-to-GDP ratio in developing countries between 1970 and 2003

	Average *log GDP*, 1970 - 1974	Average *log GDP*, 1999 - 2003	Average net FDI inflows as percentage of GDP, 1970 – 1974	Average net FDI inflows as percentage of GDP, 1999 – 2003
Argentina	26.05	27.08	0.22	2.96
Brazil	25.42	26.88	1.22	4.19
Costa Rica	22.34	23.52	2.41	3.30
Ecuador	22.60	23.54	4.69	5.14
Dom. Republic	22.41	23.72	2.48	4.73
Mexico	25.86	26.32	0.83	2.78
Chile	23.81	25.07	-1.02	6.33
Peru	24.12	24.72	-0.07	2.70
Venezuela	25.02	25.47	-0.58	2.67
Columbia	24.12	25.17	0.38	2.51
Sri Lanka	22.21	23.52	0.02	1.14
India	25.48	26.90	0.06	0.65
Pakistan	23.61	25.04	0.08	0.72
Indonesia	24.10	25.77	0.79	-1.69
Malaysia	23.36	25.24	2.86	3.10
Korea	25.45[a]	27.00	0.15[a]	1.12
Philippines	24.13	25.08	0.01	1.62
Thailand	23.79	25.57	0.81	2.64
Singapore	23.33[b]	25.22	5.98[b]	15.05
Cameroon	21.90	22.96	0.58	0.83
Zambia	21.65	21.94	-2.38	3.09
Côte d'Ivoire	22.38	23.08	1.44	2.09
Nigeria	23.80	24.50	2.14	2.44
Tunisia	22.36	23.73	1.37	2.75
Kenya	22.02	23.09	0.61	0.49
Morocco	23.21	24.29	0.20	4.00
Ghana	21.71	22.37	1.16	2.18
Egypt	23.78	25.35	0.01	0.79

Source: WDI 2005, own calculations. [a] Average over the period 1976 - 1980, [b] Average over the period 1972 – 1976.

Table 2

Cointegration test results

Country	Engle-Granger		ECM		Johansen		Model
	ADF - Statistic	1% (5%) critical values [a]	ECM - t-statistic	1 % (5%) critical values [b]	λ-trace statistic	1% (5%) critical values[c]	
Argentina	-2.51 [0]	-4.81 (-4.07)	-1.65	-4.60 (-3.84)	21.63 [1]	30.45 (25.32)	A
Brazil	-3.10 [2]	-4.81 (-4.07)	-0.89	-4.60 (-3.84)	23.41 [2]	30.45 (25.32)	A
Costa Rica	-2.68 [1]	-4.81 (-4.07)	-3.23	-4.60 (-3.84)	18.64 [1]	30.45 (25.32)	A
Ecuador	**-6.01** [0]**	**-4.81 (-4.07)**	**-4.02***	**-4.60 (-3.84)**	**58.32** [1]**	**30.45 (25.32)**	**A**
Dom. Republic	-3.81 [0]	-4.81 (-4.07)	-2.62	-4.60 (-3.84)	24.93 [1]	30.45 (25.32)	A
Mexico	**-4.76** [1]**	**-4.24 (-3.52)**	-2.38	-4.06 (-3.32)	**24.46** [1]**	**20.04 (15.41)**	**B**
Chile	-2.25 [0]	-4.81 (-4.07)	-2.85	-4.60 (-3.84)	18.70 [1]	30.45 (25.32)	A
Peru	-1.69 [0]	-4.24 (-3.52)	-2.55	-4.06 (-3.32)	7.38 [1]	20.04 (15.41)	B
Venezuela	**-3.54* [0]**	**-4.24 (-3.52)**	-2.87	-4.06 (-3.32)	**18.57* [1]**	**20.04 (15.41)**	**B**
Columbia	-1.88 [1]	-4.81 (-4.07)	-2.27	-4.60 (-3.84)	12.60 [2]	30.45 (25.32)	A
Sri Lanka	**-4.61* [5]**	**-4.81 (-4.07)**	**-4.62****	**-4.60 (-3.84)**	**28.48* [1]**	**30.45 (25.32)**	**A**
India	-3.63 [0]	-4.81 (-4.07)	-3.32	-4.60 (-3.84)	23.95 [1]	30.45 (25.32)	A
Pakistan	-2.76 [0]	-4.24 (-3.52)	0.39	-4.06 (-3.32)	14.11 [1]	20.04 (15.41)	B
Indonesia	-2.27 [1]	-4.81 (-4.07)	-1.38	-4.60 (-3.84)	16.35 [3]	30.45 (25.32)	A
Malaysia	-3.08 [1]	-4.81 (-4.07)	-2.25	-4.60 (-3.84)	20.12 [1]	30.45 (25.32)	A
Korea	-2.26 [1]	-4.31 (-3.56)	-1.57	-4.13 (-3.35)	12.54 [2]	20.04 (15.41)	B
Philippines	-1.78 [0]	-4.24 (-3.52)	-2.18	-4.06 (-3.32)	10.27 [1]	20.04 (15.41)	B
Thailand	-2.36 [1]	-4.81 (-4.07)	-1.67	-4.60 (-3.84)	13.89 [2]	30.45 (25.32)	A
Singapore	-2.41 [0]	-4.24 (-3.52)	-1.87	-4.06 (-3.32)	14.23 [1]	20.04 (15.41)	B
Cameroon	-1.32 [1]	-4.81 (-4.07)	-2.85	-4.60 (-3.84)	20.83 [1]	30.45 (25.32)	A
Zambia	-2.52 [0]	-4.81 (-4.07)	-2.27	-4.60 (-3.84)	25.28 [1]	30.45 (25.32)	A
Côte d'Ivoire	-3.21 [0]	-4.24 (-3.52)	-2.46	-4.06 (-3.32)	10.51 [1]	20.04 (15.41)	B
Nigeria	-3.57 [3]	-4.81 (-4.07)	-1.84	-4.60 (-3.84)	11.41 [1]	30.45 (25.32)	A
Tunisia	-4.02 [0]	-4.81 (-4.07)	-3.02	-4.60 (-3.84)	18.29 [3]	30.45 (25.32)	A
Kenya	-1.66 [1]	-4.81 (-4.07)	-0.52	-4.60 (-3.84)	14.94 [1]	30.45 (25.32)	A
Morocco	-1.94 [1]	-4.81 (-4.07)	-2.86	-4.60 (-3.84)	20.99 [2]	30.45 (25.32)	A
Ghana	-2.20 [0]	-4.81 (-4.07)	-1.05	-4.60 (-3.84)	21.50 [2]	30.45 (25.32)	A
Egypt	-3.12 [0]	-4.81 (-4.07)	-3.32	-4.60 (-3.84)	17.88 [1]	30.45 (25.32)	A

Numbers in brackets indicate the number of lags. The trace statistics are adjusted for degrees of freedom used in estimation. [a] Critical values from MacKinnon (1991). [b] Critical values from Ericsson and MacKinnon (2002). [c] Critical values from Osterwald-Lenum (1992). Model A: an intercept and a linear trend are included and the trend is restricted to enter the cointegrating vector (intercept and trend). Model B: allows for a trend in the data but not on the cointegrating vector (intercept, no trend). * (**) indicate a rejection of the null of no cointegration at the 5% (1%) level.

Table 3

Gregory-Hansen cointegration test results

Country	Level shift model (C)			Slope change model (C/T)			Regime shift model (C/S)		
	ADF*-statistic	1% (5%) Critical values	Break year	ADF*-statistic	1% (5%) Critical values	Break year	ADF*-statistic	1% (5%) Critical values	Break year
Argentina	-2.18 [0]	-5.13 (-4.61)	1985	-3.94 [1]	-5.45 (-4.99)	1975	-3.16 [0]	-5.47 (-4.95)	1982
Brazil	-3.08 [0]	-5.13 (-4.61)	1981	-2.65 [0]	-5.45 (-4.99)	1981	-3.43 [0]	-5.47 (-4.95)	1981
Costa Rica	-2.17 [0]	-5.13 (-4.61)	1981	-4.86 [1]	-5.45 (-4.99)	1981	-3.91 [0]	-5.47 (-4.95)	1980
Dom.Republic	-2.41 [0]	-5.13 (-4.61)	1980	-3.26 [0]	-5.45 (-4.99)	1981	-2.77 [0]	-5.47 (-4.95)	1980
Chile	-2.48 [0]	-5.13 (-4.61)	1982	-2.82 [0]	-5.45 (-4.99)	1980	-2.76 [0]	-5.47 (-4.95)	1974
Peru	-1.97 [0]	-5.13 (-4.61)	1982	-3.86 [1]	-5.45 (-4.99)	1986	-1.91 [0]	-5.47 (-4.95)	1982
Columbia	-3.16 [1]	-5.13 (-4.61)	1983	-1.98 [0]	-5.45 (-4.99)	1978	-3.00 [1]	-5.47 (-4.95)	1983
India	-2.77 [0]	-5.13 (-4.61)	1978	-4.81 [0]	-5.45 (-4.99)	1977	-3.17 [0]	-5.47 (-4.95)	1981
Pakistan	-2.78 [0]	-5.13 (-4.61)	1980	-3.29 [0]	-5.45 (-4.99)	1981	-2.80 [0]	-5.47 (-4.95)	1980
Indonesia	-2.19 [0]	-5.13 (-4.61)	1983	-4.34 [1]	-5.45 (-4.99)	1975	-2.49 [0]	-5.47 (-4.95)	1983
Malaysia	-2.47 [0]	-5.13 (-4.61)	1994	-3.37 [1]	-5.45 (-4.99)	1982	-2.76 [0]	-5.47 (-4.95)	1994
Korea	-2.47 [0]	-5.13 (-4.61)	1988	-2.90 [0]	-5.45 (-4.99)	1986	-4.11 [1]	-5.47 (-4.95)	1991
Philippines	-1.68 [0]	-5.13 (-4.61)	1974	-3.25 [1]	-5.45 (-4.99)	1974	-1.69 [0]	-5.47 (-4.95)	1974
Thailand	-1.88 [0]	-5.13 (-4.61)	1976	-2.50 [1]	-5.45 (-4.99)	1983	-2.05 [0]	-5.47 (-4.95)	1981
Singapore	-2.22 [0]	-5.13 (-4.61)	1974	-0.98 [0]	-5.45 (-4.99)	1981	-3.03 [0]	-5.47 (-4.95)	1989
Cameroon	-3.40 [0]	-5.13 (-4.61)	1982	-3.18 [5]	-5.45 (-4.99)	1974	-4.61 [0]	-5.47 (-4.95)	1982
Zambia	-3.31 [0]	-5.13 (-4.61)	1998	-3.38 [1]	-5.45 (-4.99)	1998	-3.66 [0]	-5.47 (-4.95)	1998
Côte d'Ivoire	-4.31 [0]	-5.13 (-4.61)	1975	-4.54 [0]	-5.45 (-4.99)	1975	-4.08 [0]	-5.47 (-4.95)	1975
Nigeria	-1.33 [0]	-5.13 (-4.61)	1983	**-5.61** [0]**	**-5.45 (-4.99)**	**1981**	-1.44 [0]	-5.47 (-4.95)	1983
Tunisia	-3.30 [0]	-5.13 (-4.61)	1983	**-5.10* [0]**	**-5.45 (-4.99)**	**1976**	-3.29 [0]	-5.47 (-4.95)	1983
Kenya	-2.48 [0]	-5.13 (-4.61)	1974	-3.98 [0]	-5.45 (-4.99)	1975	-2.50 [0]	-5.47 (-4.95)	1974
Morocco	-2.79 [2]	-5.13 (-4.61)	1975	-2.80 [1]	-5.45 (-4.99)	1976	-3.73 [1]	-5.47 (-4.95)	1981
Ghana	-2.32 [0]	-5.13 (-4.61)	1980	-3.75 [1]	-5.45 (-4.99)	1978	-2.07 [0]	-5.47 (-4.95)	1981
Egypt	-2.18 [0]	-5.13 (-4.61)	1978	**-5.63** [0]**	**-5.45 (-4.99)**	1979	-2.76 [0]	-5.47 (-4.95)	1976

Numbers in brackets indicate the number of lags. The lag length was determined using the t-sig method. * (**) indicate a rejection of the null of no cointegration at the 5% (1%) level. Critical values are taken from Gregory and Hansen (1996).

Following Gregory and Hansen (1996), we computed the ADF statistics for each breakpoint in the interval $0.15T$ – $0.85T$.

Table 4

Long-run relationship: Phillips and Loretan (1991) nonlinear least squares

Equation (10): Dependent variable: $LogGDP_t$												
Country	Coefficients (λ_1) / Independent variable				Diagnostic tests							
	$\hat{c}_1$	$\hat{\delta}_1 / t$	$\hat{\beta}_1 / (FDI_t / GDP_t)$	$\hat{\mu}$ / $\varphi_{t\tau}$	$\bar{R}^2$	SE	$LM(1)$	$LM(3)$	$Arch(1)$	$Arch(3)$	JB
Ecuador	22.80**	0.030**	-0.036*		0.99	0.03	0.119	0.265	0.417	0.389	2.708
	(468.7)	(10.28)	(-2.87)				(0.733)	(0.850)	(0.524)	(0.762)	(0.258)
Sri Lanka	22.16**	0.041**	0.098**		0.99	0.01	0.269	1.841	0.010	0.059	2.708
	(880.5)	(30.81)	(4.09)				(0.610)	(0.178)	(0.921)	(0.943)	(0.258)
Nigeria	23.77**	0.023**	0.049*	-0.199*	0.92	0.05	1.341	2.006	0.102	0.522	1.588
	(631.1)	(4.105)	(2.43)	(-2.64)			(0.266)	(0.167)	(0.753)	(0.672)	(0.452)
Tunisia	22.35**	0.040**	0.027*	0.065*	0.99	0.02	2.467	1.289	0.756	0.391	0.086
	(764.0)	(42.66)	(2.66)	(2.83)			(0.131)	(0.306)	(0.392)	(0.761)	(0.957)
Egypt	23.65**	0.049**	0.047**	0.123**	0.99	0.02	0.941	0.562	0.464	0.445	4.319
	(876.1)	(30.22)	(3.74)	(2.885)			(0.346)	(0.648)	(0.501)	(0.723)	(0.115)
Equation (11): Dependent variable: FDI_t / GDP_t											
Country	Coefficients (λ_2) / Independent variable				Diagnostic tests						
	$\hat{c}_2$	$\hat{\beta}_2 / Log\ GDP_t$			$\bar{R}^2$	SE	$LM(1)$	$LM(3)$	$Arch(1)$	$Arch(3)$	JB
Mexico	-155.8**	6.037**			0.83	0.41	0.058	0.151	0.000	0.239	1.528
	(-10.58)	(10.69)					(0.813)	(0.928)	(0.99)	(0.868)	(0.466)
Venezuela	-2.75.9**	10.93**			0.59	0.29	0.005	0.383	0.876	0.369	2.981
	(-5.549)	(5.572)					(0.942)	(0.766)	(0.357)	(0.776)	(0.225)

* (**) denote the 5% (1%) level of significance.
For Ecuador, Tunisia, Venezuela [Mexico, Sri Lanka] (Nigeria) the Phillips-Loretan equation was estimated with up to one [two] (three) leads and lags and with up to one lag of the equilibrium error. For Egypt the Phillips-Loretan equation was estimated with up to two leads and lags and with up to two lags of the equilibrium error.

Table 5

Long-run causality tests

Country	t-Value of α_1 ($LogGDP_t$)	t-Value of α_2 (FDI_t/GDP_t)	Conclusion
Ecuador	-4.16**	0.50	$FDI_t / GDP_t \rightarrow LogGDP_t$
Sri Lanka	-3.41**	3.53**	$FDI_t / GDP_t \leftrightarrow LogGDP_t$
Nigeria	-2.93**	2.58*	$FDI_t / GDP_t \leftrightarrow LogGDP_t$
Tunisia	-3.84**	4.53**	$FDI_t / GDP_t \leftrightarrow LogGDP_t$
Egypt	-3.51**	5.19**	$FDI_t / GDP_t \leftrightarrow LogGDP_t$
Mexico	0.52	-4.11**	$FDI_t / GDP_t \leftarrow LogGDP_t$
Venezuela	1.13	-3.80**	$FDI_t / GDP_t \leftarrow LogGDP_t$

* (**) denote the 5% (1%) level of significance. Corresponding variables which were tested for weak exogeneity are in parentheses. $FDI_t / GDP_t \rightarrow LogGDP_t$ stands for "long-run causality runs from FDI_t / GDP_t to $LogGDP_t$" and vice versa.

Table 6

Short-run causality

Country	Dependent variable: $\Delta LogGDP_t$		Dependent variable: $\Delta(FDI_t / GDP_t)$		Conclusion
	Coefficients of $\sum\Delta(FDI_t / GDP_t)$	Coefficients of $\sum\Delta LogGDP_t$	Coefficients of $\sum\Delta LogGDP_t$	Coefficients of $\sum\Delta(FDI_t / GDP_t)$	
Argentina	—	0.324^{+} (3.49)[1]	—	—	No causality
Brazil	—	0.491** (10.05)[1]	—	0.624** (13.33)[1]	No causality
Costa Rica	—	0.197* (3.95)[2]	—	-2.307* (5.32)[1]	No causality
Dom. Rep.	—	0.364* (4.66)[1]	—	—	No causality
Chile	**0.008* (3.92)[1]**	0.392* (7.56)[1]	—	-0.439* (7.06)[1]	$FDI_t / GDP_t \rightarrow LogGDP_t$
Peru	**0.012^{+} (3.46)[1]**	0.299^{+} (3.22)[1]	—	—	$FDI_t / GDP_t \rightarrow LogGDP_t$
Columbia	**-0.011* (7.49)[2]**	0.440** (8.40)[1]	—	-0.588** (14.99)[2]	$FDI_t / GDP_t \rightarrow LogGDP_t$
India	**0.086^{+} (3.23)[2]**	—	—	—	$FDI_t / GDP_t \rightarrow LogGDP_t$
Pakistan	—	—	—	—	No causality exists
Indonesia	—	0.286** (12.41)[1]	**12.48** (9.65)[2]**	-0.460* (7.24)[2]	$FDI_t / GDP_t \leftarrow LogGDP_t$
Malaysia	—	—	—	—	No causality
Korea	—	—	—	1.011** (13.42)[2]	No causality
Philippines	—	0.526** (11.54)[1]	—	-0.415* (5.01)[1]	No causality
Thailand	**0.008^{+} (3.07)[1]**	0.397** (9.98)[1]	**8.647* (6.81)[2]**	—	$FDI_t / GDP_t \leftrightarrow LogGDP_t$
Singapore	—	—	—	—	No causality
Cameroon	—	—	—	-0.705** (30.55)[1]	No causality
Zambia	**-0.003^{+} (3.70)[2]**	—	—	—	$FDI_t / GDP_t \rightarrow LogGDP_t$
C. d'Ivoire	—	0.436* (7.12)[1]	—	-0.318^{+} (3.55)[1]	No causality
Kenya	**0.015^{+} (4.03)[1]**	0.548** (12.43)[2]	—	-1.04* (7.24)[2]	$FDI_t / GDP_t \rightarrow LogGDP_t$
Morocco	—	-0.510** (10.6)[1]	—	-0.901** (81.01)[1]	No causality
Ghana	—	—	**8.260* (4.14)[1]**	—	$FDI_t / GDP_t \leftarrow LogGDP_t$
Ecuador	—	—	—	-0.146^{+} (3.32)[1]	No short-run causality
Sri Lanka	**-0.161* (3.74)[2]**	—	—	—	$FDI_t / GDP_t \rightarrow LogGDP_t$
Nigeria	—	—	—	—	No short-run causality
Tunisia	—	—	**9.049^{+} (3.10)[2]**	—	$FDI_t / GDP_t \leftarrow LogGDP_t$
Egypt	**-0.008* (7.16)[2]**	0.76** (27.74)[1]	—	—	$FDI_t / GDP_t \rightarrow LogGDP_t$
Mexico	—	—	—	—	No short-run causality
Venezuela	—	—	—	—	No short-run causality

Numbers in brackets indicate the number of lags. The numbers in parentheses are *F*-statistics for the joint significance of the lagged variables. $^{+}$ [*] (**) denote the 10% [5%] (1%) level of significance. $FDI_t / GDP_t \rightarrow LogGDP_t$ stands for "short-run causality runs from FDI_t / GDP_t to $LogGDP_t$" and vice versa. For Chile an impulse dummy for 1982, for Indonesia, Malaysia, Korea, and Thailand an impulse dummy for 1998 was needed to achieve a normal distribution of the residuals. For Ecuador an impulse dummy for 1976 was included in the $\Delta(FDI_t / GDP_t)$ equation, but not in the $\Delta LogGDP_t$ equation. The impulse dummy variables are one in the corresponding years and zero otherwise. For Ghana a linear trend was included in the $\Delta LogGDP_t$ equation (but not in the $\Delta(FDI_t / GDP_t)$ equation).

Figure 1

Cross-plots of the long-run impact of FDI on GDP and development indicators

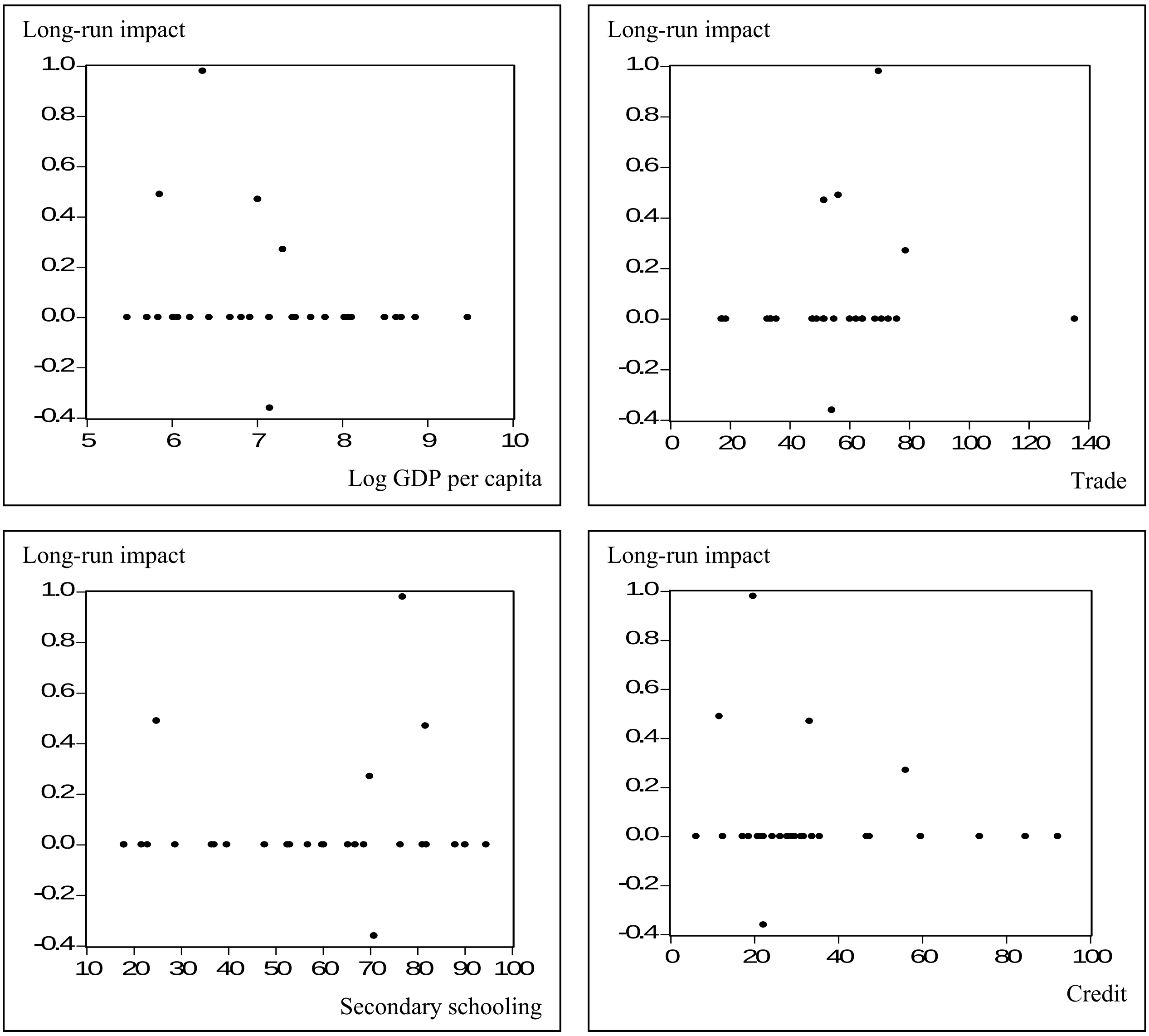

Figure 2

Cross-plots of the short-run impact of FDI on GDP and development indicators