

Herzer, Dierk

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# The Long-Run Relationship between Outward FDI and Total Factor Productivity: Evidence for Developing Countries

Dierk Herzer<sup>\*</sup>

*University of Wuppertal, Germany*

## **Abstract**

This paper examines the long-run relationship between outward foreign direct investment (FDI) and total factor productivity for a sample of 33 developing countries over the period 1980-2005. Using panel cointegration techniques, we find that: (i) outward FDI has, on average, a positive long-run effect on total factor productivity in developing countries, (ii) increased factor productivity is both consequence and a cause of increased outward FDI, and (iii) there are large differences in the long-run effects of outward FDI on total factor productivity across countries. Cross-sectional regressions indicate that these cross-country differences in the productivity effects of outward FDI are significantly negatively related to cross-country differences in labor market regulation, whereas there is no statistically significant association between the productivity effects of outward FDI and the level of human capital, the level of financial development, or the degree of trade openness in the home country.

*JEL classification:* F21; O11; F23; C23

*Keywords:* outward FDI; total factor productivity; developing countries, panel cointegration

## **1. Introduction**

Foreign direct investment (FDI) outflows from developing countries have grown faster in recent decades than those from developed economies. The share of developing countries in total world FDI outflows increased more than thirty-fold from approximately 0.5% in the early 1970s to about 16% in 2008. In 2008, FDI outflows from developing countries reached almost 300 billion US\$, which is more than three times the value of world FDI outflows in 1970. The stock of outward FDI by developing country firms was \$2.36 trillion in 2008, accounting for about 15% of the world outward FDI stock in that year.<sup>1</sup>

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<sup>\*</sup> *Correspondence Address:* Dierk Herzer, Schumpeter School of Business and Economics, University of Wuppertal, Gaußstr. 20, 42119 Wuppertal, Germany. Email: herzer@wiwi.uni-wuppertal.de.

<sup>1</sup> The figures are based on data from the UNCTAD FDI database (<http://stats.unctad.org/FDI/ReportFolders/reportFolders.aspx>).

Despite the enormous increase in FDI outflows from developing countries, there is surprisingly little evidence on the potential economic consequences of outward FDI for these countries. The empirical literature regarding outward FDI consists mainly of firm and industry-level studies on the effects of outward FDI on employment, exports, investment, and productivity in developed economies, while only a few such micro studies are available on developing countries. Macroeconomic studies of the overall impact of outward FDI on the home developing countries have not yet been performed.

This paper attempts to fill this gap by examining the long-run relationship in developing countries between outward FDI and total factor productivity as a key to economic growth. Specifically, we aim to answer the following questions:

- (1) Does outward FDI reduce or increase total factor productivity in developing countries on average over the long run?
- (2) Is outward FDI an exogenous influence on total factor productivity or does the causality run in both directions?
- (3) Are the productivity effects of outward FDI constant or do they differ across countries?

To answer these questions, we apply panel cointegration techniques to a sample of 33 developing countries over the period 1980-2005. Panel cointegration estimators are robust under cointegration to a variety of estimation problems that often plague empirical work, including omitted variables, endogeneity, and measurement errors (see, e.g., Banerjee, 1999; Baltagi and Kao, 2000; Pedroni, 2007). Moreover, panel cointegration methods can be implemented with shorter data spans than their time-series counterparts. Our main results are: (1) outward FDI has, on average, a positive long-run effect on total factor productivity in developing countries, (2) increased factor productivity is both consequence and a cause of increased outward FDI, and (3) there are large differences in the long-run effects of outward FDI on total factor productivity across countries.

Given this latter finding, a further question arises:

- (4) How can the cross-country differences in the long-run productivity effects be explained?

As an additional contribution, we attempt to answer this question by examining whether the observed cross-country differences in the long-run effects of outward FDI are linked to differences in country characteristics, such as the levels of human capital, financial market development, trade openness, and labor market regulation. Using cross-sectional analysis, we find that the cross-country differences in the productivity effects of outward FDI are negatively and significantly related to cross-country differences in labor market regulation, whereas there is no statistically significant association between the productivity effects of outward FDI and the level of human capital, the level of financial development, and the degree of trade openness in the home country.

The remainder of this paper is composed of four sections. In Section 2, we discuss the theoretical background and related empirical literature. Section 3 sets out the basic empirical model and describes the data. The econometric implementation and the estimation results are presented in Section 4, while Section 5 concludes.

## 2. Theoretical background and related empirical literature

### 2.1. Theoretical background

As a background for the discussion of the potential theoretical effects of outward FDI on domestic productivity in developing countries, it is useful to briefly describe the characteristics of outward FDI from these countries and to contrast them with the pattern of outward investment from developed economies.

Table 1

Outward FDI from developing countries by sector and destination, percentage shares

Outward FDI stock by sector, 2006 <sup>a</sup>		
	Developing countries	Developed countries
Primary	3.4	7.9
Manufacturing	9.4	28.5
Services	83.7	61.7
Unspecified	3.5	1.9
Outward FDI flows by destination, 2000-2004 <sup>b</sup>		
	Destination	
	Developing countries	Developed countries
Developing countries <sup>c</sup>	74.2	25.8
Developed countries	15.9	78.3

Notes: <sup>a</sup> Source: UNCTAD (2008), own calculations; <sup>b</sup> averages over the period 2000-2004, source: UNCTAD (2006), own calculations; <sup>c</sup> FDI from developing and transition economies excluding offshore financial centers to developing and transition economies

A striking feature of outward FDI from developing countries is the dominance of the service sector, as shown in Table 1. In 2006, 83.7% of the stock of their outward FDI was in services, such as trade, finance, and business activities (see, e.g., UNCTAD, 2006); in developed countries, the corresponding share was 61.7%. The second largest sector is manufacturing, which accounted in 2006 for 9.4% of the outward FDI stock of developing countries and 28.5% of the outward FDI stock of developed countries. Outward FDI in the primary sector, in contrast, is of minor importance. In 2006, only 3.4% of the stock of outward FDI from developing countries was in this

sector, while the corresponding figure for developed-country outward FDI was 7.9%. Perhaps most striking, however, is that developing countries heavily invest in other developing countries, mostly from the same region (see, e.g., UNCTAD, 2006). In the period 2000-2004, approximately 75% of FDI outflows from developing countries went to the developing world. The opposite is true for developed countries that tend to predominantly direct their investment towards other developed countries. Finally, it is also worth mentioning that, although the number of developing-country multinationals on the top-100 list has increased in recent years, most of them are relatively small compared to multinationals from developed countries (see, e.g., UNCTAD, 2006).

With this background, we begin our discussion of the potential productivity effects of outward FDI in developing countries by considering possible interactions between domestic and foreign activities of multinational firms. To represent these interactions, suppose that the production function of a multinational firm is given by  $Q=f(D, F, \theta_d, \theta_f)$ , where  $Q$  is the representative firm's worldwide output,  $D$  is domestic input,  $F$  is foreign input,  $\theta_d$  is a vector of factors that influence domestic production (such as domestic productivity), and  $\theta_f$  represents factors that influence foreign production. Assuming that domestic input production is a function of domestic capital,  $D(K_d)$ , and that foreign input production is a function of foreign capital,  $F(K_f)$ , the first-order condition that characterizes the firm's profit-maximizing choice of domestic inputs is

$$\frac{\partial Q [D(K_d), F(K_f), \theta_d, \theta_f]}{\partial D(K_d)} = \lambda, \quad (1)$$

where  $\lambda$  is total input cost — the firm's cost of capital. From Equation (1) it can be seen that domestic and foreign production (or investment) of the multinational firm can be related either through the cost of capital, and thus through the financial side of the firm, if  $\lambda$  is somehow a function of  $F$ , or through the production process, if  $\partial^2 Q [D(K_d), F(K_f), \theta_d, \theta_f] / \partial D(K_d) \partial F(K_f)$  is nonzero (see also Desai et al., 2005).

Interactions between foreign and domestic activities operating through the financial side of the firm occur in a situation where fixed investments in different locations compete for funds due to costly external financing, as discussed in more detail by Stevens and Lipsey (1992). In such a scenario, the decision to invest scarce resources abroad inevitably reduces the likelihood of concurrent investments at home, implying that each dollar of outward FDI displaces a dollar of domestic investment (see also, e.g., Feldstein, 1994; Desai et al., 2005; Herzer and Schrooten, 2008). This substitution of domestic for foreign investment, in turn, is likely to also reduce domestic productivity. In particular, when the investments abroad come at the expense of

investments necessary to sustain productivity at home (such as new machinery, worker training, and research and development (R&D)), outward FDI may reduce the domestic productivity of the investing firm in the long run.

Some studies, however, suggest that the situation of fixed resources appears to be rather atypical for multinational firms, at least for developed-country multinationals. Desai et al. (2004), for example, analyze how US multinationals capitalize affiliates around the world and find that US multinational affiliates substitute internal borrowing for costly external finance stemming from adverse capital market conditions. Similarly, Desai et al. (2008) show that US parents provide affiliates with additional equity to finance profitable investment opportunities during currency crises. However, these findings cannot necessarily be generalized to developing-country multinationals. Since multinationals from developing countries are relatively small compared to their developed-country counterparts, they often do not have internal capital markets and must therefore rely on external financing to finance their investment projects. In addition, domestic financial markets are undeveloped in many developing countries. Thus, although some developing-country multinational companies have access to foreign capital markets (see, e.g., UNCTAD, 2006), they are generally more likely to face financial constraints than developed-country multinationals. The conclusion from this is that it is also more likely for developing-country than for developed-country multinationals that outward FDI leads to a reduction in domestic productivity due to potential interactions between domestic and foreign activities through the financial side.

The financial side is not the only source of possible interactions between domestic and foreign activities of multinational firms, as shown above. A second and perhaps more important source of interdependence is the production process. Because of production interdependence, outward FDI can affect domestic productivity in several ways, each of which depends on the multinational firm's investment motive and the respective investment type. In the following, we distinguish four key types of investment: horizontal FDI, vertical FDI, strategic asset-seeking FDI and resource-seeking FDI.

Horizontal or market-seeking FDI is motivated by market access and avoidance of trade frictions such as transport costs and import protection in the host country (for models of horizontal FDI see, e.g., Markusen, 1984; Horstmann and Markusen, 1987; Markusen and Venables, 1998). The decision to engage in horizontal FDI is guided by the proximity-concentration tradeoff in which proximity to the host market avoids trade costs but incurs the added fixed cost of building a second production facility. FDI of this type thus occurs when a firm decides to serve foreign markets through local production, rather than exports, and hence to produce the same product or service in multiple countries. Consequently, horizontal FDI may substitute for exports of the goods that were

previously produced in the investor's home country. This decrease in domestic export production, in turn, may be accompanied by a decrease in domestic productivity, since export intensity and firm productivity may be linked, as some studies suggest (see, e.g., Castellani, 2002; Baldwin and Gu, 2003). However, such effects occur only if the produced good is tradable. As discussed above, the overwhelming majority of outward FDI from developing countries is in services, most of which are non tradable, implying that the bulk of horizontal developing-country FDI cannot substitute for home-country exports. But even those horizontal investments that are directed to tradable sectors do not necessarily reduce domestic exports and productivity. The reason is because there is rarely a pure case of horizontal production in the sense that there is inevitably some vertical component to a firm, horizontal FDI can boost exports of intermediate goods and services from the home to the host country. For example, headquarters in the home country provide specialized services to foreign affiliates (such as R&D, design, marketing, finance, strategic management) even if the same final goods are produced in both the home and foreign country (see, e.g., Kokko, 2006). Thus, in general terms, multinational firms combine home production with foreign production to increase their productivity and hence competitiveness both internationally and domestically (see, e.g., Herzer, 2008; Desai et al., 2009). Furthermore, in the long run, horizontal FDI may allow the firm to raise its competitiveness through access to new markets or successful penetration of existing markets, thereby additionally increasing domestic productivity.

The horizontal motive is the most important for outward FDI from developing countries, followed by the vertical motive (see, e.g., UNCTAD, 2006). Vertical or efficiency-seeking FDI is driven by international factor price differences (for models of vertical FDI see, e.g., Helpman, 1984; Helpman and Krugman, 1985). It takes place when a firm fragments its production process internationally, locating each stage of production in the country where it can be done at the lowest cost. Such relocations reduce domestic production, at least in the short run (as with horizontal FDI). However, in the longer run, vertical investment may allow the firm to import cheaper intermediate inputs from foreign affiliates and/or to produce a greater volume of final goods abroad at lower cost, thereby stimulating exports of intermediate goods used by foreign affiliates (see, e.g., Herzer, 2008). The new structure of the production chain may thus be associated with increased efficiency, and, as a result, the firm may be able to improve its competitive position, thus raising its domestic productivity over the long run (see, e.g., Kokko, 2006). On the other hand, if the firm is not able to adjust over the longer term to the reduction in domestic production by failing to raise its competitiveness (e.g., due to labor market rigidities), both vertical and horizontal FDI will substitute foreign activities for domestic activities over the long run, which may also lead to a long-term decrease in domestic productivity (see, e.g., Bitzer and Görg, 2009).

The third most important motive for outward FDI from developing countries is the strategic asset-seeking motive (see, e.g., UNCTAD, 2006). As the name implies, firms undertake strategic asset-seeking FDI to acquire assets that are not available in their own country. Strategic asset-seeking FDI is made, for example, when investors attempt to gain access to internationally recognized brand names and local distribution networks in order to strengthen their international competitive position. Strategic asset-seeking FDI also occurs in the form of technology-sourcing FDI when firms attempt to gain access to foreign technology by either purchasing foreign firms or establishing R&D facilities in “foreign centers of excellence” (for models of technology-sourcing FDI, see, e.g., Neven and Siotis, 1996; Fosfuri and Motta, 1999; Bjorvatn and Eckel, 2006). If foreign affiliates then acquire new knowledge in terms of technological know-how, management techniques, knowledge of consumer tastes, etc., this knowledge can be transferred back to the parent company, thus increasing domestic productivity in the long term (see, e.g. Van Pottelsberghe and Lichtenberg, 2001). However, since the ability to absorb knowledge from abroad depends on the absorptive capacity of the investing firm, firms with low levels of technological capacity are likely to be unable to effectively access and exploit foreign knowledge through outward FDI. The potential productivity gains for outward investors from developing countries may therefore be smaller than is the case for their developed-country counterparts. Another consideration related to knowledge spillovers through outward FDI is that a substantial percentage of outward FDI from developing markets goes to other developing countries (as shown in Table 1). This South-South outward FDI may not generate significant knowledge spillovers because the bulk of FDI is not located in clusters of specific technological expertise. On the other hand, narrower technological gaps between home and host firms may facilitate absorption of technological knowledge, implying that South-South FDI may generate more spillovers than South-North FDI (see, e.g., Moran, 2008). In addition, developing-country multinationals have a greater propensity to establish linkages with local firms than do their counterparts from developed countries (see, e.g., UNCTAD, 2006), which in turn enables them to more deeply integrate into the host economies, and this deeper integration could be particularly beneficial in terms of reverse knowledge and technology flows back to the home country (see, e.g., Globerman and Shapiro, 2008).

Finally, the least important, yet still significant, motive for investors from developing countries is the resource-seeking motive (see, e.g., UNCTAD, 2006). Resource-seeking FDI occurs when firms identify specific host country locations as attractive source of natural resources at the lowest cost. Such FDI is usually associated with exports of resource-based products from the host country and should improve the productivity of domestic production which uses the imported resources as low cost, high quality inputs.



An important point is that outward FDI may not only affect the productivity of the investing firms, but also that of the economy as a whole through productivity spillovers to local firms (see, e.g., Blomström and Kokko, 1998). For example, local firms may improve their productivity by copying technologies used by domestic multinationals, or domestic producers may benefit from the knowledge and expertise of the outward-investing firms through labor turnover. Moreover, the increased competition between international firms and their domestically oriented counterparts may force the latter to use their existing resources more efficiently. Outward-investing firms, due to the increased productivity, may be also able to provide higher quality inputs at lower prices to local producers. In addition, if outward FDI allows the investing firms to grow larger than would be possible with production in just one country, both the investing companies and their local suppliers may benefit from economies of scale. Outward FDI may thus enable domestic suppliers to move down their learning curves and, therefore, to realize substantial productivity gains.

Since, however, FDI may act as an important vehicle for the transfer of technological and managerial know-how, it is likely to increase the competitiveness of the host economy as well. This may lead to reductions in domestic output and productivity when domestic consumers prefer the foreign competitors. Furthermore, the increased competitiveness may allow domestic firms in the host country to challenge the foreign firms and thereby to capture market shares from the foreign affiliates of the home country's multinationals. Outward FDI may therefore enable competitors in the host country to attract demand away from the home country firms, forcing them to reduce their production and to move up their average cost curve, resulting in productivity losses in the home country. In addition, outward FDI can reduce domestic capital accumulation and thus domestic productivity when outward investors claim scarce domestic resources, such as domestic financial capital, that could otherwise have been used by domestic investors for investment in their home country (see, e.g., UNCTAD, 2006).

Thus, the net effect of outward FDI on aggregate productivity in developing countries is theoretically ambiguous and must be determined empirically. Although there are no empirical investigations on this overall macroeconomic effect on developing countries, some studies do exist on the firm- and industry-level effects of outward FDI on domestic productivity for both developing and developed countries. Also, there is some evidence of cross-border R&D spillovers through outward FDI at the macro-level for developed economies.

## 2.2. Available evidence

Van Pottelsberghe and Lichtenberg (2001) use country-level macro data for a panel of 13 developed countries over the period 1971-1990 to examine whether technology-sourcing FDI affects domestic productivity through foreign R&D spillovers. They find a positive long-run relationship between the foreign R&D capital stock weighted by outward FDI and domestic total factor productivity, implying that outward FDI into R&D-intensive countries indeed has beneficial effects upon home-country productivity by transferring technological knowledge from the host country. However, Bitzer and Kerekes (2008) reach a different conclusion. Their findings, based on industry-level data for 17 OECD countries between 1973 and 2000, suggest that the interaction between foreign R&D capital and outward FDI is negatively associated with domestic productivity in non-G7 countries; for the G7 the evidence of R&D spillovers through outward FDI is not significant. Both studies investigate only whether outward FDI into major R&D-performing countries acts as a channel for R&D spillovers, thus neglecting all other potential productivity effects of outward FDI.

Braconier et al. (2001), in contrast, investigate both the effect of the outward-FDI-weighted foreign R&D capital stock and the effect of “pure” outward FDI on domestic productivity. Using firm- and industry-level data for Sweden over the period 1978-1994, they find neither evidence of FDI-related R&D spillovers, nor any correlation between outward FDI per se and domestic productivity for Sweden. These results differ from those of Driffield et al. (2009). In an industry study for the UK covering the period 1978-1994, the authors distinguish between outward FDI in high-cost, high-R&D-intensive and outward FDI in low-cost, low-R&D-intensive countries. They find that both types of FDI generate productivity growth in the UK, suggesting that technology-sourcing and efficiency-seeking FDI increase domestic productivity. A similar result is obtained by Driffield and Chiang (2009), who investigate the effects of outward FDI from Taiwan to China. Based on industry data for 1995-2005, they report a positive association between outward FDI to China and labor productivity in Taiwan. Given the fact that labor costs in Taiwan are significantly above those in China, the authors conclude that this productivity effect is due to vertical (efficiency-seeking) FDI. Vahter and Masso (2007), on the other hand, use firm-level panel data from Estonia between 1995 and 2002 to examine the effects of outward FDI on total factor productivity of the investing firms and the rest of the industry. They find that outward FDI is positively related to the productivity of the parent companies, whereas there is no robust evidence of productivity spillovers to other firms. Since the overwhelming majority of FDI by Estonian firms is horizontal (market-seeking) (see, e.g., Masso et al., 2008), the positive productivity effects of Estonian outward FDI appear to be primarily associated with this type of FDI.

In another study, Kimura and Kiyota (2006) analyze Japanese firm-level data for the period 1994-2000. One of their findings is that outward FDI increases firm productivity. More specifically, their results suggest that firms engaging in outward FDI experience, on average, productivity growth 1.8% higher than domestic firms not engaging in outward FDI. Hijzen et al. (2006a), however, criticize this study for failing to control for the endogeneity bias that arises when more productive firms self-select into investing abroad. To deal with this endogeneity problem, they apply matching and difference-in-differences analysis to data of Japanese firms for the period 1995-2002. The evidence in their study suggests that the effect of outward FDI on Japanese firm productivity is not significant.

Propensity score matching and difference-in-difference techniques are also used in other studies. Barba Navaretti and Castellani (2004) apply these methods to Italian firm-level data for the period 1973-1991. They find that multinational firms have higher total factor productivity growth after investing abroad than does a counterfactual of national firms. Kleinert and Toubal (2007), in an analysis of German firm-level data for years 1997-2003, find no significant effect from the establishment of a foreign affiliate on firm productivity growth. Hijzen et al. (2006b), using French firm-level data between 1984 and 2002, report that firms that invest in developed countries increase their productivity, while firms that invest in developing countries experience no productivity effects, which could suggest that productivity effects of outward FDI are primarily associated with horizontal investment rather than vertical investment. Barba Navaretti et al. (2009) obtain the same result in a sample of French firms in the period 1993-2000; for Italy, however, they find exactly the opposite pattern: firms that invest in developing countries experience an increase in total factor productivity, whereas FDI into developed countries has no productivity effects.

Finally, Bitzer and Görg (2009) examine the effect of outward FDI on domestic total factor productivity using industry-level panel data for 17 OECD countries over the period 1973-2001. Their results suggest that outward FDI has, on average, a negative effect on total factor productivity, but that there are large differences across countries. Outward FDI has the largest negative effect on total factor productivity in South Korea — the only developing country in the sample. In France, Japan, Poland, Sweden, the Czech Republic, the UK, and the US, in contrast, increased outward FDI is associated with higher total factor productivity.

Given the mixed results, perhaps the only conclusions that can safely be drawn from these studies are that outward FDI can have positive, as well as negative, effects on domestic productivity, that the domestic productivity effects of outward FDI do not necessarily depend upon the investment motive, and that the effects of outward FDI can differ significantly from country to country. The latter may apply in particular to developing countries, which differ widely in terms of

country size, income level, economic structure, natural resources, technological capabilities, trade openness, government policies, and other characteristics. Unfortunately, the studies do not provide any information on how outward FDI could affect aggregate productivity in developing countries on average over the long run.

### 3. Empirical model and data

The analysis will examine the long-run relationship between outward FDI and total factor productivity in developing countries. In this section, we present the basic empirical model, discuss some econometric issues, and describe the data.

#### 3.1. Empirical specification and econometric issues

We assume that the correct specification of the functional form of the long-run relationship between total factor productivity and outward FDI is given by

$$\log(TFP_{it}) = a_i + \delta_i t + b \log(OFDI_{it}) + \varepsilon_{it}, \quad (2)$$

where  $i = 1, 2, \dots, N$  is the country index,  $t = 1, 2, \dots, T$  is the time index,  $\log(TFP_{it})$  represents the log of total factor productivity, and  $\log(OFDI_{it})$  is the log of outward FDI. Following Bitzer and Görg (2009), we use outward FDI stocks rather than outward FDI flows, because stocks, due to the accumulation of flows, may more effectively capture long-run effects. The size of the long-run effect of outward FDI on total factor productivity is measured by the coefficient  $b$ , which can be interpreted as the long-run elasticity of total factor productivity with respect to outward FDI. Finally, any country-specific omitted factors which are relatively stable in the long run or evolve smoothly over time are captured by country-specific fixed effects,  $a_i$ , and country-specific time trends,  $\delta_i t$ .

Equation (2) assumes a long-run bivariate relationship between permanent movements in the log level of outward FDI and permanent movements in the log level of total factor productivity. Necessary conditions for this assumption to hold (and thus for our model to be a correct description of the data) are that both the individual time series for the log of total factor productivity and the individual time series for the log of outward FDI are nonstationary or, more specifically, integrated of the same order and that  $\log(TFP_{it})$  and  $\log(OFDI_{it})$  form a cointegrated pair. A regression consisting of two cointegrated variables has a stationary error term,  $\varepsilon_{it}$ , in turn implying that no

relevant integrated variables are omitted; any omitted nonstationary variable that is part of the cointegrating relationship would enter the error term, thereby producing nonstationary residuals and thus leading to a failure to detect cointegration.

Another assumption inherent in Equation (2) is that total factor productivity is endogenous in the sense that, in the long run, changes in outward FDI cause changes in total factor productivity. However, although the existence of cointegration implies long-run Granger-causality in at least one direction, long-run causality may also run from total factor productivity to outward FDI. The rationale is that recent theoretical work on firm heterogeneity and FDI suggests that the establishment or acquisition of foreign affiliates involves additional costs of overcoming legal, cultural and social barriers, so that only firms above a certain productivity threshold can cope with these fixed costs and thus engage in outward FDI (see, e.g., Helpman et al., 2004; Aw and Lee, 2008). That is, only the most productive firms self-select into investing abroad.<sup>2</sup> Since an increase in aggregate productivity is generally associated with an increase in average firm productivity and, consequently, with an increase in the number of firms reaching the critical productivity level necessary for FDI, a macroeconomic implication of heterogeneous-firm models is that the aggregate amount of outward FDI should increase as total factor productivity increases. On the other hand, given that total factor productivity growth is generally associated with domestic output growth, higher demand, and hence better profit opportunities for domestic investment, an increase in total factor productivity can also lead to a reallocation of scarce funds to more profitable domestic investment opportunities in place of less profitable outward investment. Consequently, increased factor productivity may be both the cause of reduced and the cause of increased outward FDI activity. The empirical implication is that it is not only crucial to examine the time-series properties of the variables and to test whether the variables are cointegrated, but it is also important to deal with this endogeneity problem and to investigate the direction of causality.

Further, an econometric issue is the potential cross-country heterogeneity in the relationship between outward FDI and total factor productivity. As discussed in Section 2.2, individual country studies based on firm- and industry-level data tend to find different results for different countries. In particular, the study by Bitzer and Görg (2009) suggests that the productivity effects of outward FDI are not constant across countries. Thus, we face a dilemma regarding the optimal estimation strategy. On the one hand, efficiency gains from the pooling of observations over the cross-

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<sup>2</sup> Models of firm heterogeneity typically predict a productivity ranking in which foreign investing firms are the most productive, followed by exporters and non-exporters. This is confirmed by empirical evidence which shows that multinationals are the most productive among the three types of firms, see, e.g., Head and Ries (2003) for Japan, Girma et al. (2004) for Ireland, Arnold and Hussinger (2005) for Germany, and Girma et al. (2005) for the UK.

sectional units can be achieved when the individual slope coefficients are the same,  $b_i = b$ . On the other hand, pooled within-dimension estimators produce inconsistent and potentially misleading point estimates of the sample mean of the heterogeneous cointegrating vectors when the true slope coefficients are heterogeneous (see, e.g., Pesaran and Smith, 1995). Although a comparative study by Baltagi and Griffin (1997) concludes that the efficiency gains from pooling more than offset the biases due to individual country heterogeneity, we try to solve this dilemma by using both homogeneous (within-dimension-based) and heterogeneous (between-dimension-based) estimators. The latter allows us, in addition, to explicitly analyze the potential heterogeneity in the effects of outward FDI across countries.

A final econometric issue is the potential cross-sectional dependence among the variables. In particular, the total factor productivities are likely to be cross-sectionally dependent, as suggested by both theoretical models and empirical studies. Neoclassical growth models, for example, implicitly assume that all countries have access to the same global stock of knowledge, while many endogenous growth models predict that knowledge diffuses across national boundaries through channels such as trade and FDI. Since the total factor productivity of a country depends upon the available (local and foreign) technological knowledge, it follows that factor productivities are globally interdependent. This conclusion is supported by several studies showing that a country's productivity depends not only upon its own stock of knowledge but also upon the knowledge stocks, and thus the total factor productivities, of its trading partners (see, e.g., Coe and Helpman, 1995; Van Pottelsberghe and Lichtenberg, 2001; Luintel and Khan, 2004). Given that standard panel unit root and cointegration tests may be biased in the presence of such cross-sectional dependence, we also use recent advances in panel data econometrics to account for this issue.

### *3.2. Data and descriptive statistics*

We now describe the data used in the empirical analysis. Total factor productivity is defined in the usual way as  $TFP = Y / [K^{(1-\alpha)} L^\alpha]$ , where  $Y$  is output,  $K$  denotes the capital stock,  $L$  stands for labor input,  $1-\alpha$  is the capital share of income, and  $\alpha$  is the labor share of income. We assume a constant  $\alpha$  of 0.6667, which can be justified as follows: First, it is common practice in the literature to assume (or use) a constant labor share of two-thirds; and second, the evidence of Gollin (2002) suggests that the labor share is, in fact, approximately constant across time and space with a value of about two-thirds. Output is measured by real GDP (in 2000 US\$), while  $L$  is represented by the labor force (i.e., the number of persons of working age, defined as 15-64 years). Without question, a better measure of labor input would be employment times average hours, but reliable data on

employment and hours worked are generally not available for developing countries over a long enough period of time. Therefore, we follow the common practice in developing country studies and use the labor force as our measure of labor input (see, e.g., Abu-Qarn and Abu-Bader, 2007). Unfortunately, official capital stock data are also not available for most developing countries. We therefore construct the physical capital stock from real investment data (gross capital formation in 2000 US\$) using the perpetual inventory equation  $K_t = I_t + (1 - \delta)K_{t-1}$ , where  $I_t$  is investment and  $\delta$  is the depreciation rate. Consistent with the literature, we set the initial value of the capital stock equal to  $K_0 = I_0 / (g + \delta)$ ,<sup>3</sup> where  $I_0$  is the value of the investment series the first year it is available, and  $g$  is the average growth rate of the investment series between the first year with available data and the first year of the estimation period (see, e.g., Caselli, 2005). In cases where the time interval between the first year with available data and the first year of the estimation period is less than five years, we follow Luintel and Khan (2004) and use the average growth rate of investment over the estimation period. As is standard in the literature, a depreciation rate of 6% is assumed. All data used to calculate total factor productivity are from the World Development Indicators (WDI) 2008 CD-Rom.

The data on outward FDI stocks are obtained from the UNCTAD FDI database (<http://www.unctad.org/templates/Page.asp?IntItemID=3277&lang=1>). UNCTAD (2008) defines FDI as an investment involving a long-term relationship and reflecting lasting interest and control of a resident entity in one economy in an enterprise resident in an economy other than that of the foreign direct investor. The stock of FDI is defined as the value of the share of the foreign enterprise capital and reserves (including retained profits) attributable to the parent enterprise plus the net indebtedness of affiliates to the parent enterprise. Since UNCTAD reports outward FDI stocks as shares of GDP, we multiply the outward FDI-to-GDP ratio by real GDP (from the WDI) to construct real outward FDI stocks (in 2000 US\$). Given that the UNCTAD data start in 1980 while the WDI 2008 data end in 2005, the empirical analysis covers the period 1980-2005.

Our sample includes all countries for which data in this period are available. Of these countries, four are in North Africa (Algeria, Egypt, Morocco, and Tunisia), nine are in sub-Saharan Africa (Benin, Botswana, Burkina Faso, Gabon, Kenya, Mali, Senegal, South Africa, and Swaziland), three are in Central America (Costa Rica, Mexico, and Panama), nine are in South America (Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Paraguay, Peru, and Venezuela), five are in East Asia (Hong Kong, Indonesia, Malaysia, South Korea, and Thailand), and three are

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<sup>3</sup> The rationale for this choice is that  $I / (g + \delta)$  is the expression for the capital stock in the steady state of the Solow model.

in South Asia and the Middle East (India, Jordan, and Pakistan). Thus, our sample consists of 33 countries from all developing regions of the world and should therefore be reasonably representative of developing countries.

Table 2 lists the countries along with the average values for  $TFP_{it}$  and  $OFDI_{it}$  over the period of observation. As can be seen, there are considerable differences in the values of these parameters across countries. Hong Kong is the country with the highest productivity level, followed by Argentina, South Korea, and Mexico, while Burkina Faso ranks at the bottom of the productivity scale. Total factor productivity in Hong Kong exceeds the productivity index of Burkina Faso by almost 20-fold. Perhaps even more striking are the cross-country differences in the outward FDI stocks. Brazil's FDI stock is almost 20 000 times larger than the FDI stock of Burkina Faso. Brazil heads the list of foreign investors, followed by Hong Kong, South Africa, and Argentina. The countries with the lowest stocks of FDI are (in ascending order) Burkina Faso, Benin, Bolivia, and Tunisia. Altogether, it appears that countries with higher factor productivities tend to have larger outward FDI stocks, while those with lower productivities correspondingly have smaller FDI stocks, suggesting a positive relationship between these two variables.

Table 2  
Countries and country summary statistics

	Average of $TFP$	Average of $OFDI$		Average of $TFP$	Average of $OFDI$
Algeria	224.102	177255922.3	Jordan	269.823	114670074
Argentina	595.990	15086190138	Kenya	92.284	80298706
Benin	88.621	3253730.583	Malaysia	305.768	2319106268
Bolivia	176.481	9672677.323	Mali	55.858	20317464.95
Botswana	296.739	535659528.9	Mexico	499.795	5971847601
Brazil	368.927	56519048833	Morocco	192.061	234835656.8
Burkina Faso	51.087	2841393.57	Pakistan	117.824	235902533.6
Chile	427.631	1051618514	Panama	386.538	5362104852
Colombia	247.698	869019532.9	Paraguay	182.837	160216629.5
Costa Rica	408.307	59784139.85	Peru	251.160	148649580.6
Ecuador	173.295	27312641.43	Senegal	155.525	50770812.4
Egypt	213.389	342765973.8	South Africa	346.997	16664398977
Gabon	336.409	155474639.7	South Korea	514.345	4173795953
Hong Kong	962.892	28538838397	Swaziland	207.713	44361876.87
India	81.456	361307750.7	Thailand	147.820	385186217.3
Indonesia	103.942	368874659.6	Tunisia	235.176	17800529.28
			Venezuela	494.961	1853046953

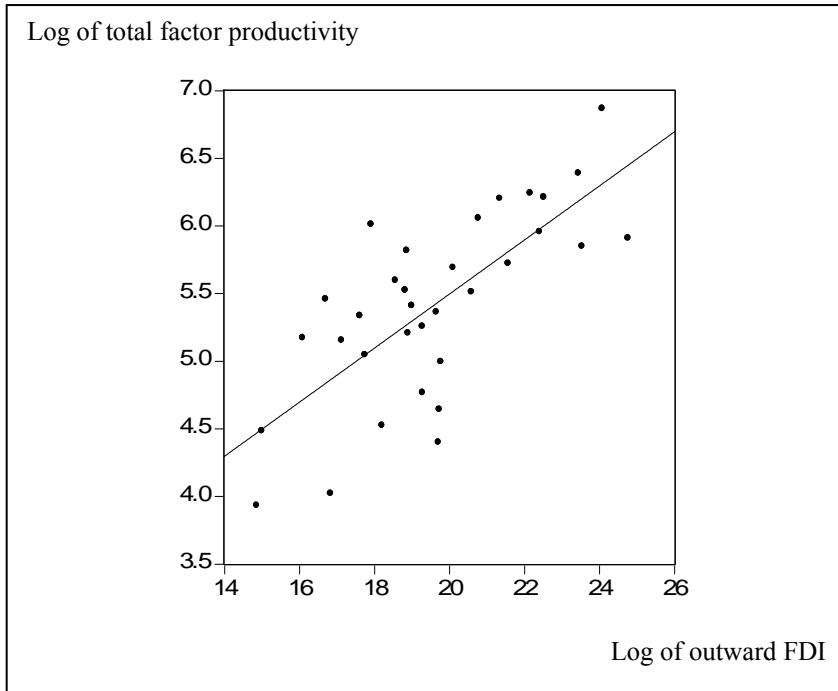
Figure 1 illustrates this cross-country relationship graphically. It shows the scatter plot of the average log of total factor productivity versus the average log of outward FDI for the 33 countries in our sample along with the regression line. The slope of the regression line is positive (and statistically significant with a  $t$  statistic of 5.94), indicating that increased outward FDI is indeed



positively associated with increased factor productivity. In the next section, we examine this relationship in more detail using panel cointegration and causality techniques.

Figure 1

Scatter plot of the log of total factor productivity versus the log of outward FDI



#### 4. Empirical analysis

The pre-tests for unit-roots and cointegration, which are reported in the Appendix, suggest that the variables are nonstationary and cointegrated, as assumed in Equation (2). In this section, we provide estimates of the cointegrating relationship between outward FDI and total factor productivity and test the robustness of the estimates. We also investigate the direction of causality between the two variables, examine the degree of heterogeneity in the effects of outward FDI on total factor productivity across countries, and search for explanations for the cross-country heterogeneity.

##### 4.1. Long-run relationship

In order to estimate the long-run elasticity of total factor productivity with respect to outward FDI, we use the dynamic ordinary least squares (DOLS) estimator. This estimator is asymptotically unbiased and normally distributed even in the presence of endogenous regressors

(see, e.g., Stock and Watson, 1993), thus allowing us to control for the potential endogeneity of outward FDI discussed in Section 3.1. Furthermore, the DOLS estimator performs well in finite samples compared with other cointegration estimators (such as the fully modified estimator) both in time-series and panel data (see, e.g., Stock and Watson, 1993; Kao and Chiang, 2000). The within-dimension-based DOLS model used in this paper and following Kao and Chiang (2000) is

$$\log(TFP_{it}) = a_i + \delta_i t + b \log(OFDI_{it}) + \sum_{j=-k}^k \Phi_{ij} \Delta \log(OFDI_{it-j}) + \varepsilon_{it}, \quad (3)$$

where  $\Phi_{ij}$  are coefficients of current, lead and lag differences, which account for possible serial correlation and endogeneity of the regressor(s), thus yielding unbiased estimates.

The results of this estimation procedure are presented in the first column of Table 3 where, for brevity, we report only the estimated  $b$  coefficients. The estimated coefficient is highly significant and positive. More precisely, the elasticity of total factor productivity with respect to outward FDI is estimated to be 0.024, implying that, in the long run, a 1% increase in the outward FDI stock is associated with an increase in total factor productivity by 0.024%. From this it can be concluded that developing countries benefit in general or on average from outward FDI due to the increased productivity of the investing companies and associated productivity spillovers to local firms.

Table 3

Estimates of the long-run effect,  $b$ , of outward FDI on total factor productivity

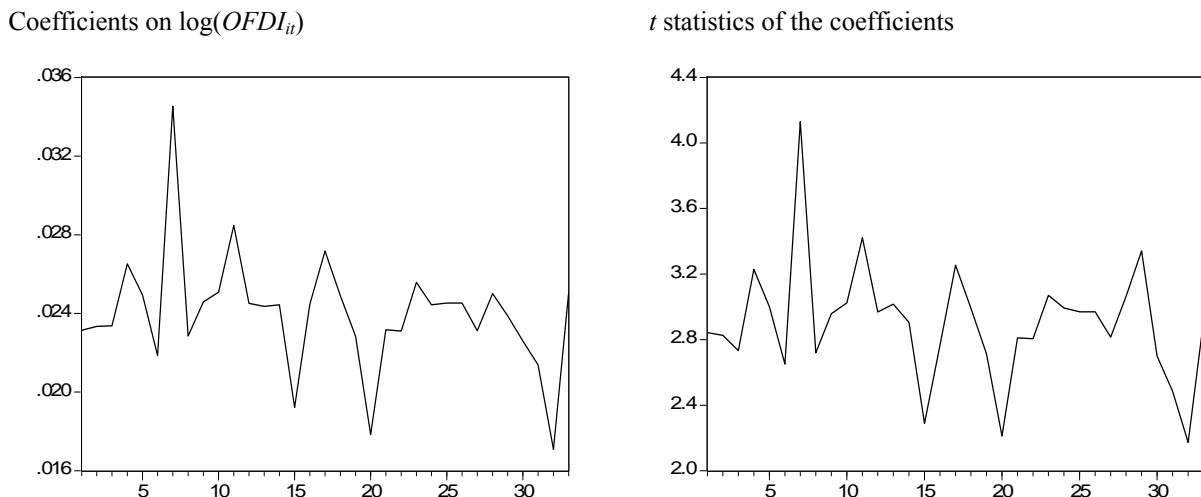
Within-dimension DOLS estimator (Kao and Chiang, 2000)	DOLS mean group estimator (Pedroni, 2001)	CCE mean group estimator (Pesaran, 2006)
0.024** (2.97)	0.014** (3.42)	0.050** (4.61)

Notes: The dependent variable is  $\log(TFP_{it})$ .  $t$ -statistics in parenthesis. \*\* indicate significance at the 1% level. The DOLS regressions were estimated with one lead and one lag.

To assess the robustness of this conclusion, we perform several sensitivity checks. First, we investigate whether the positive relationship between outward FDI and total factor productivity is robust to alternative estimation methods. Specifically, a potential problem with the above estimation procedure could be that it assumes a homogeneous  $b$ , which is likely to be empirically incorrect (as we will show in Section 4.3). To allow the slope coefficients to vary across countries, we use the between-dimension, group-mean panel DOLS estimator suggested by Pedroni (2001). This estimator involves estimating separate DOLS regressions for each country and averaging the long-run coefficients,  $\hat{b} = N^{-1} \sum_{i=1}^N \hat{b}_i$ . The  $t$  statistic for the average is the sum of the individual  $t$  statistics divided by the root of the number of cross-sectional units,  $t_{\hat{b}} = \sum_{i=1}^N t_{\hat{b}_i} / \sqrt{N}$ . We present the DOLS

group-mean point estimate of the effect of outward FDI on total factor productivity in the second column of Table 3. Since, however, the DOLS estimates may be biased in the presence of cross-sectional dependence, we also report (in the third column) the result of the common correlated effects (CCE) mean group estimator suggested by Pesaran (2006). This estimator allows for cross-sectional dependencies that potentially arise from multiple unobserved common factors and is the simple average of the individual common correlated effects estimators given by Equation (A.6) in the Appendix. As can be seen, all three estimators produce similar results, suggesting that the positive relationship between outward FDI and total factor productivity is neither due to potentially restrictive homogeneity assumptions nor to possible cross-sectional dependence. However, given the relatively short time dimension of our data, the mean group results (which are based on individual time-series regressions) should be interpreted with caution. In addition, the CCE mean group estimator is intended for the case in which the regressors are exogenous, so that we lose the ability to account for the likely endogeneity of outward FDI. Therefore, we continue our robustness analysis with the pooled within-dimension panel DOLS estimator. As noted in Section 3.1, there is evidence to suggest that the efficiency gains from pooling are likely to more than offset the potential biases due to individual heterogeneity (see, e.g., Baltagi and Griffin, 1997).

Figure 2  
DOLS estimation with single country excluded from the sample



We re-estimate the DOLS regression excluding one country at a time from the sample to verify that the positive effect of outward FDI is not due to individual outliers. The sequentially estimated long-run coefficients and their  $t$  statistics are presented in Figure 2. As they are relatively

stable between 0.017 and 0.035 and always significant at the 5% level, we conclude that the positive productivity effect is not the result of individual outliers.

Next, we examine whether the positive long-run relationship between outward FDI and total factor productivity is due to sample-selection bias. Sample-selection bias occurs when the selected sample is not random and thus not representative. A potential problem with our sample could be that it includes only 6 low-income countries (Benin, Burkina Faso, Kenya, Mali, Pakistan, Senegal), while the rest of the countries are classified by the World Bank (WDI 2008 CD-Rom) as middle- or high income countries. Another possible problem is the distribution of outward FDI among the countries. In fact, our sample is dominated by 6 countries (Argentina, Brazil, Hong Kong, Mexico, Panama, South Africa) that invest more than the sample average. We therefore re-estimate the DOLS regression for four subsamples: low-income countries, middle- and high-income countries, countries with outward FDI above the sample average, and countries with outward FDI below the sample average. The resulting coefficients are listed in Table 4. Regardless of which subsample is used, the long-run relationship between outward FDI and total factor productivity remains positive and significant at least at the 5% level. From this, it can be concluded that the positive coefficient on  $\log(OFDI_{it})$  is not due to sample-selection bias. Clearly, it would be desirable to also assess whether there are significant differences in the effects of outward FDI on total factor productivity between low-income and middle- to high-income countries or between countries with low levels of outward FDI and those with high levels of outward FDI. However, the small sample sizes do not allow statistically meaningful comparisons.

Table 4

DOLS estimates for subsamples

	$\log(OFDI_{it})$	Number of countries in the subsample
Low-income countries	0.079* (2.16)	6
Middle- and high-income countries	0.028** (3.08)	27
Countries with outward FDI above the sample average	0.056* (2.19)	6
Countries with outward FDI below the sample average	0.021* (2.38)	27

Notes: \*\* (\*) indicate significance at the 1% (5%) level. *t*-statistics in parentheses.

Finally, we examine whether the estimated coefficient on outward FDI is sensitive to the model specification. To this end, we combine the definition of total factor productivity,  $TFP = Y/[K^{(1-\alpha)}L^\alpha]$ , with Equation (1) to obtain (after some rearrangements)

$$\log\left(\frac{Y_{it}}{L_{it}}\right) = a_{it} + \delta_i t + (1-\alpha)\log\left(\frac{K_{it}}{L_{it}}\right) + b\log(OFDI_{it}) + \varepsilon_{it}, \quad (4)$$

where  $(Y_{it}/L_{it})$  is labor productivity,  $(K_{it}/L_{it})$  is capital per worker, and  $\alpha$  is (as before) the labor share of income, which was assumed to be 0.6667. Since the elasticity of labor productivity with respect to outward FDI,  $b$ , is equal to the elasticity of total factor productivity with respect to outward FDI, Equation (4) allows us to test whether there are systematic biases in the calculated factor productivity and the estimated coefficient on outward FDI. If there are no systematic biases, Equation (4) should, on the one hand, produce approximately the same outward FDI coefficient as Equation (2) — and thus a  $b$  value of about 0.024 — and, on the other hand, a coefficient on log capital per worker of about  $(1 - 0.6667 =) 0.3333$ . The DOLS estimates for these two parameters are given in Table 5. As can be seen, the coefficient on log capital per worker is indeed close to 1/3 and the estimated  $b$  coefficient is close to 0.024. In addition, the last row of Table 5 shows that simple Wald tests do not reject the restrictions that  $(1 - \hat{\alpha}) = 0.3333$  and  $b = 0.024$ . Thus, it can be concluded that the positive productivity effect of outward FDI is robust to different estimation techniques, potential outliers, sample selection, and the specification of the empirical model.

Table 5

DOLS estimates of the coefficients on log capital per worker,  $(1 - \alpha)$ , and log outward FDI,  $b$

	$\log(K_{it}/L_{it})$	$\log(OFDI_{it})$
	0.368**	0.019**
	(14.14)	(2.99)
Wald tests	Restriction: $(1 - \hat{\alpha}) = 0.3333$	Restriction: $\hat{b} = 0.24$
$\chi^2(1)$	1.81	0.52
[ $p$ -values]	[0.18]	[0.47]

Notes: The dependent variable is  $\log(Y_{it}/L_{it})$ . \*\* indicate significance at the 1% level.  $t$ -statistics in parentheses. The DOLS regression was estimated with one lead and one lag. The number of degrees of freedom  $\nu$  in the  $\chi^2(\nu)$  tests correspond to the number of restrictions.

#### 4.2. Long-run causality

Recent theories on firm heterogeneity and FDI suggest that only those firms with productivities above a certain threshold find it profitable to locate production or other activities abroad (see, e.g., Helpman et al., 2004; Aw and Lee, 2008). Since it is reasonable to assume that an increase in aggregate productivity is associated with an increase in the number of firms reaching this threshold, it follows that economy-wide productivity gain should lead to increased FDI outflows. Consequently, the positive coefficient on  $\log(OFDI_{it})$  does not necessarily reflect a causal effect of outward investment on total factor productivity, but causality may also run from  $\log(TFP_{it})$  to  $\log(OFDI_{it})$ .

To test the direction of causality, we use a two-step procedure. In the first step, we employ the DOLS estimate of the long-run relationship to construct the disequilibrium term

$$ec_{it} = \log(TFP_{it}) - [\hat{a}_i + \hat{\delta}_i t + 0.024 \log(OFDI_{it})]. \quad (5)$$

In the second step, we estimate the error correction model (ECM)

$$\begin{aligned} \Delta \log(TFP_{it}) &= c_{1i} + a_1 ec_{it-1} + \sum_{j=1}^k \varphi_{11ij} \Delta \log(TFP_{it-j}) + \sum_{j=1}^k \varphi_{12ij} \Delta \log(OFDI_{it-j}) \\ \Delta \log(OFDI_{it}) &= c_{2i} + a_2 ec_{it-1} + \sum_{j=1}^k \varphi_{21ij} \Delta \log(TFP_{it-j}) + \sum_{j=1}^k \varphi_{22ij} \Delta \log(OFDI_{it-j}) \end{aligned}, \quad (6)$$

where the error-correction term,  $ec_{it-1}$ , represents the error in, or deviation from, the equilibrium, while the adjustment coefficients  $a_1$  and  $a_2$  capture how  $\log(TFP_{it})$  and  $\log(OFDI_{it})$  respond to deviations from the equilibrium relationship. From the Granger representation theorem, we know that at least one of the adjustment coefficients must be non-zero if a long-run relationship between the variables is to hold. A significant error-correction term also suggests long-run Granger causality and thus long-run endogeneity (Hall and Milne, 1994), whereas a non-significant adjustment coefficient implies long-run Granger non causality from the independent to the dependent variable(s) as well as weak exogeneity. In the following, we test for weak exogeneity of total factor productivity and outward FDI, and thus for long-run Granger non causality between  $\log(TFP_{it})$  and  $\log(OFDI_{it})$ , by first successively eliminating the insignificant short-run run dynamics with the lowest  $t$  values. We then test the significance of the adjustment coefficients. In doing so, we reduce the number of parameters (according to Hendry's general-to-specific methodology) and thereby we increase the precision of the weak exogeneity tests on the  $a$ -coefficients. Since all variables in the model, including  $ec_{it-1}$ , are stationary (because the level variables are integrated of order 1 and cointegrated), a conventional likelihood ratio test can be used to test the null hypothesis of weak exogeneity,  $H_0 : a_{1,2} = 0$ .

However, the above model assumes that the adjustment coefficients as well as the short- and long-term effects are the same for all countries. To account for heterogeneous long-run elasticities, we replace the error-correction term given by Equation (5) with the residuals from the individual DOLS long-run relationships:

$$ec_{it} = \log(TFP_{it}) - [\hat{a}_i + \hat{\delta}_i t + \hat{b}_i \log(OFDI_{it})]. \quad (7)$$

Although this model allows the log-run coefficients to differ across countries, the short-run parameters and adjustment coefficients are still restricted to be the same. Because this homogeneity assumption may also be violated, we finally allow for complete heterogeneity by estimating the ECM with the error-correction term given by Equation (7) separately for each country. More specifically, we proceed as follows: We successively eliminate the insignificant short-run dynamics

from the ECM and compute the  $p$  values of the likelihood ratio test for the null hypothesis of weak exogeneity for each country individually; the panel weak exogeneity (or long-run Granger non causality) test is then conducted using the Fisher statistic proposed by Madalla and Wu (1999):

$$\lambda = -2 \sum_i^N \log(p_i), \quad (8)$$

where  $p_i$  is the  $p$  value of the of the likelihood test for country  $i$ . The Fisher statistic is distributed as  $\chi^2$  with  $2 \times N$  degrees of freedom.

Table 6

Weak exogeneity tests / long-run causality tests

Assumptions		Weak exogeneity of $\log(TFP_{it})$ (Significance of $\alpha_1$ )	Weak exogeneity of $\log(OFDI_{it})$ (Significance of $\alpha_2$ )
Homogeneous long-run coefficients, short-run parameters and adjustment coefficients	$\chi^2(1)$ ( $p$ -values)	186.01 (0.000)	10.58 (0.001)
Heterogeneous long-run coefficients, homogeneous short-run parameters and adjustment coefficients	$\chi^2(1)$ ( $p$ -values)	91.14 (0.000)	6.12 (0.013)
Heterogeneous long-run coefficients, short-run parameters and adjustment coefficients	Fisher statistics ( $p$ -values)	369.19 (0.000)	136.60 (0.000)

*Notes:* The number of degrees of freedom  $\nu$  in the standard  $\chi^2(\nu)$  tests correspond to the number of zero restrictions. The Fisher statistic is distributed as  $\chi^2$  with  $2 \times N$  degrees of freedom. It has a critical value of 95.62 at the 1% level. The number of lags was determined by the general-to-specific procedure with a maximum of three lags.

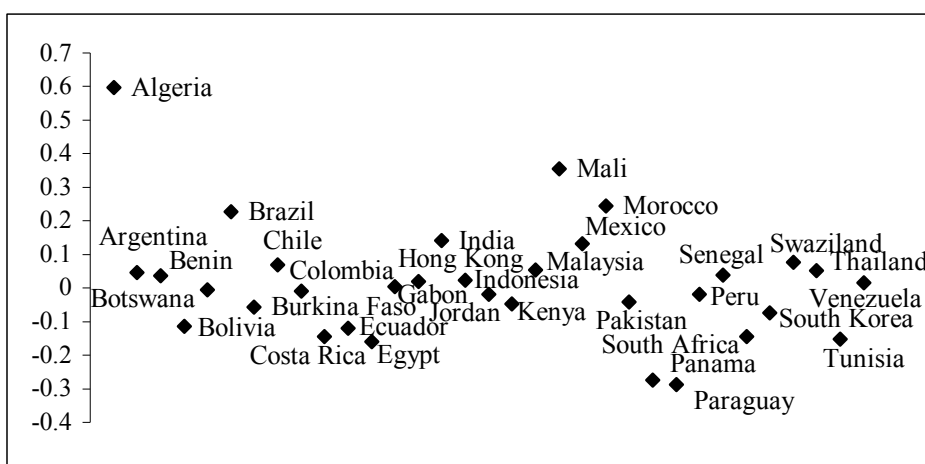
Table 6 presents the results. Regardless which model is employed, the null hypothesis of weak exogeneity is rejected for both  $\log(TFP_{it})$  and  $\log(OFDI_{it})$  at least at the 5% level. From this it can be concluded that the statistical long-run causality is bidirectional, suggesting that increased factor productivity is both a consequence and a cause of increased outward FDI. Thus, the evidence in this paper also supports the macroeconomic implication of heterogeneous-firm models that outward FDI tends to increase when a country's aggregate productivity increases.

### 4.3. Cross-country heterogeneity

The results reported so far indicate that outward FDI has, on average, a positive long-run effect on total factor productivity in developing countries (and vice versa). This finding for the sample as a whole does not imply, however, that outward FDI exerts positive productivity effects in each individual country. Figure 3 plots the individual country estimates of the coefficient on  $\log(OFDI_{it})$  from the group-mean panel DOLS estimator. As previously noted, these estimates must be interpreted with caution given the short sample period. What can be safely concluded, however, is that there is considerable heterogeneity in the effects of outward FDI on total

productivity across countries. The coefficients range from -0.29 in Paraguay to 0.60 in Algeria, indicating that, although the long-run effect of outward FDI on total factor productivity is positive in general or on average in developing countries, outward FDI does not have a positive long-run effect on total factor productivity in all countries. More specifically, we find that in 17 countries (Algeria, Argentina, Benin, Brazil, Chile, Gabon, Hong Kong, India, Indonesia, Malaysia, Mali, Mexico, Morocco, Senegal, Swaziland, Thailand, and Venezuela), an increase in outward FDI is associated with an increase in total factor productivity, while in 16 cases (Bolivia, Botswana, Burkina Faso, Colombia, Costa Rica, Ecuador, Egypt, Jordan, Kenya, Pakistan, Panama, Paraguay, Peru, South Africa, South Korea, and Tunisia), an increase in outward FDI is associated with a decrease in total factor productivity. Note that the finding of a negative productivity effect in South Korea supports the study of Bitzer and Görg (2009) who also report a negative coefficient on outward FDI for that country (as discussed in Section 2.2). But even within the country groups with negative and positive effects, the individual country estimates show considerable heterogeneity. For example, the point estimates suggest that Algeria and Mali benefit markedly from outward FDI. In contrast, in many countries, such as Botswana and Gabon, both the positive and negative effects are marginal (close to 0), whereas in some other countries, such as Paraguay and Panama, outward FDI has a strong, negative effect on total factor productivity. Altogether, the negative effects tend to be smaller, in absolute value, than the positive effects; the positive impact coefficients exceed the negative impact coefficients on average by about 20% (in absolute value).

Figure 3  
Individual country DOLS estimates





#### 4.4. Explanations for the cross-country heterogeneity

The cross-country differences in the long-run effect of outward FDI on total factor productivity pose a new question: What factors can explain this heterogeneity or, in other words, what factors determine the long-run effect of outward FDI on domestic productivity? A possible way to answer this question is to examine whether the observed pattern of the long-run effects of outward FDI can be linked to cross-country differences in human capital, financial development, trade openness, or labor market regulation.

The choice of these variables is inspired by the empirical literature on *inward* FDI and economic growth. Borensztein et al. (1998), for example, find that the effect of inward FDI on growth depends on the level of human capital in the host country. According to the study of Balasubramanyam et al. (1996), the effects of inward FDI on growth are stronger for countries that are more open to trade. Alfaro et al. (2004) find that FDI plays an important role in contributing to economic growth; however, the level of development of local financial markets is crucial for these positive effects to be realized. Finally, the results of Busse and Groizard (2008) suggest that the growth effect of inward FDI is negatively related to the level of regulation in the host country.

The reasons why these variables act as determinants of the growth effects of *inward* FDI are similar to the reasons why the variables could be important in explaining the cross-country differences in the aggregate productivity effects of *outward* FDI. They are as follows. Outward FDI can have beneficial effects on home country productivity by transferring technological knowledge from the host country, as discussed above. The ability to absorb foreign knowledge and technology depends, however, on the absorptive capacity of the home country, which, in turn, is closely related to the level of human capital. That is, developing countries with low levels of human capital may be unable to make effective use of knowledge spillovers through outward FDI. Similarly, it can be argued that knowledge spillovers are typically realized only if both outward investing firms and domestic producers have the ability to invest in absorbing foreign knowledge, which may be restricted by underdeveloped local financial markets. Greater trade openness, in contrast, may promote trade between parent firms and their foreign affiliates, which increases the scope for intra-firm specialization and economies of scale, thus leading to higher efficiency of outward FDI. Finally, restrictive or costly labor market regulations may prevent both the efficient allocation of resources to foreign investing firms — which are the most productive — and the creation of linkages and spillovers to local firms. That is, the efficiency of outward FDI is reduced to the extent that labor market regulations impede the expansion of domestic multinationals and their local suppliers. Hence, it can be hypothesized that the productivity effect of outward FDI depends on the

level of human capital, the level of financial market development, the degree of trade openness, and the degree of labor market regulation.

We use the secondary school enrolment rate ( $SCHOOL_i$ ) as a proxy for human capital, the ratio of domestic credit to the private sector to GDP ( $CREDIT_i$ ) is our measure of financial development, openness is represented by the Sachs and Warner openness index ( $OPENNESS_i$ ), and labor market regulation is measured by the Heritage Foundation's labor freedom index ( $LABORFREEDOM_i$ ). Note that a higher labor freedom index indicates less regulatory constraints.<sup>4</sup> The data on schooling and financial development are from the World Development Indicators 2008 CD-Rom, the Sachs and Warner openness index is constructed on the basis of the liberalization dates provided by Sachs and Warner (1995) and Wacziarg and Welch (2008), and the labor freedom data are from <http://www.heritage.org/research/features/index/downloads.cfm>. Since the estimated productivity effects of outward FDI can be interpreted as time averages over the period 1980-2005, we also use averaged data for the above variables over that period. An exception is the labor freedom index for which data before 2005 are not available; accordingly, we are constrained to use values for that single year.<sup>5</sup> Moreover, we do not have complete data on trade openness for Panama and Swaziland, forcing us to limit our sample for the openness variable to 31 countries.

To examine the relationship between the estimated productivity effect of outward FDI and the four variables, we first regress  $\hat{b}_i$  separately on  $SCHOOL_i$ ,  $CREDIT_i$ ,  $OPENNESS_i$ ,  $LABORFREEDOM_i$  (and an intercept). Since an estimated dependent variable may introduce heteroskedasticity into the regressions, we use Newey-West heteroskedasticity and autocorrelation consistent standard errors to compute the  $t$ -statistics. The results of these regressions are reported in the columns 2 to 5 of Table 7. They show that the long-run effect of outward FDI on total factor productivity is neither significantly correlated with the level of human capital (measured by the secondary school enrolment rate) nor with the level of financial market development or trade openness. In contrast, we find a negative and significant association between the productivity effect of outward FDI and labor market regulation. (Note the sign of the coefficient on  $LABORFREEDOM_i$  is positive, given that a higher value of the labor freedom index indicates a lower level of labor market regulation.) These results do not change substantially when the four variables are included jointly in the regression, as shown in column 6.  $SCHOOL_i$ ,  $CREDIT_i$ , and  $OPENNESS_i$  are insignificant while the effect of outward FDI on total factor productivity is

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<sup>4</sup> The following four quantitative components each account for 25 percent of the labor freedom factor: minimum wage, rigidity of hours, difficulty of firing redundant employees, and cost of firing redundant workers. For more information on the labor factor, see Kane et al. (2007).

<sup>5</sup> This should not be a problem since the indices of labor freedom appear to be relatively stable over time, at least for the years 2005-2008.

significantly greater for countries with less labor market regulation. Admittedly, the coefficient on  $LABORFREEDOM_i$  is only significant at the 10% level, but this may be due to collinearity between the labor regulation variable and the other regressors;  $LABORFREEDOM_i$  is relatively highly correlated with  $SCHOOL_i$ ,  $CREDIT_i$ , and  $OPENNESS_i$ , with correlation coefficients between 0.41 and 0.59.

Table 7

Regression of the estimated long-run effects of outward FDI on indicators for human capital, financial market development, openness, and labor market regulation

Independent variables	(2)	(3)	(4)	(5)	(6)
$SCHOOL_i$	-0.00089 (-0.71)				-0.00018 (-0.79)
$CREDIT_i$		-0.00046 (-0.96)			-0.00050 (-0.67)
$OPENNESS_i$			-0.11717 (-1.02)		-0.14905 (-0.78)
$LABORFREEDOM_i$				0.00266* (2.24)	0.00398 <sup>+</sup> (1.92)
Observations	33	33	31	33	31

Notes: \* (<sup>+</sup>) indicates significance at the 5% (10%) level. Reported  $t$ -statistics (in parentheses) are based on Newey and West's heteroskedasticity and autocorrelation consistent standard errors. A higher labor freedom index ( $LABORFREEDOM_i$ ) indicates less regulatory constraints.

Without question, our sample of 33 or 31 countries is too small to draw definite conclusions regarding systematic variations in the long-run effects of outward FDI across countries. Nevertheless, despite the small sample size, our results suggest that cross-country differences in the long-run effects of outward FDI on total factor productivity can be significantly explained by cross-country differences in labor regulation, while the level of human capital and of financial development, and the degree of trade openness appear to have no statistically significant effect on the efficiency of outward FDI.

## 5. Conclusion

Numerous studies exist on the effects of outward FDI on home countries. The focus of this literature has traditionally been on outward FDI from developed economies, reflecting the fact that for a long time FDI originated almost exclusively in the developed world. However, outward FDI from developing countries has been growing significantly in both absolute and relative importance in recent years. Nevertheless, there is surprisingly little research on the home-country effects of outward FDI for these countries.

In this study, we examined the macroeconomic relationship between outward FDI and total factor productivity for developing countries, a relationship that has not yet been explored in the published literature. Our results suggest that: (1) outward FDI has, on average, a positive long-run effect on domestic total factor productivity in developing countries, (2) increased factor productivity is both a consequence and a cause of increased outward FDI, (3) there is considerable heterogeneity in the long-run effects of outward FDI on total factor productivity across countries, and (4) part of this heterogeneity can be explained by cross-country differences in labor market regulation, whereas the level of human capital, the level of financial development, and the degree of trade openness are not significantly related to cross-country variations in the domestic productivity effects of outward FDI.

In conclusion we can say that outward FDI can bring significant benefits to developing countries by increasing domestic productivity and thus economic growth. However, there are also countries for which an increase in outward FDI is associated with a decrease in total factor productivity. The central question is: How can governments in developing countries restructure their economies to make positive effects more likely and negative effects less likely? One answer provided in this paper is that the efficiency of outward FDI can potentially be increased by reducing labor market rigidities. But differences in labor market regulation are certainly not the only explanatory factor for the differences in the effects of outward FDI across countries; many other factors (such as institutions, political stability, resource endowments, etc.) are potentially important for some, but not necessarily all, countries. These and other issues on outward FDI from developing countries require further research.

## Appendix A1. Panel unit-root tests

We discuss three panel unit-root tests along with results from their application to the data in this study. Consider the following augmented Dickey-Fuller (ADF) regression where the variable of interest is observed for  $N (= 33)$  cross-sectional units and  $T (= 26)$  time periods:

$$\Delta x_{it} = z_{it}\gamma_i + \rho_i x_{it-1} + \sum_{j=1}^{k_i} \phi_{ij} \Delta x_{it-j} + \varepsilon_{it}, \quad i = 1, 2, \dots, N, \quad t = 1, 2, \dots, T, \quad (\text{A.1})$$

where  $k_i$  is the lag length,  $z_{it}$  is a vector of deterministic terms, such as fixed effects or fixed effects plus individual trends, and  $\gamma_i$  is the corresponding vector of coefficients.

The within-dimension-based panel unit-root test of Levin, Lin, and Chu (2002) (LLC) pools the autoregressive coefficient across the countries during the unit-root test and thus restrict the first-order autoregressive parameter to be the same for all countries,  $\rho_i = \rho$ . The null hypothesis is that all series contain a unit root,  $H_0 : \rho = 0$ , while the alternative hypothesis is that no series contains a unit root,  $H_1 : \rho = \rho_i < 0$ , that is, all are (trend) stationary. To conduct the LLC-test statistic, the following steps are performed. The first is to obtain the residuals,  $\hat{e}_{it}$ , from individual regressions of  $\Delta x_{it}$  on its lagged values (and on  $z_{it}$ ),  $\Delta x_{it} = \sum_{j=1}^{k_i} \theta_{1ij} \Delta x_{it-j} + z_{it}\gamma_i + e_{it}$ . Second,  $x_{it-1}$  is regressed on the lagged values of  $\Delta x_{it}$  (and on  $z_{it}$ ) to obtain  $\hat{v}_{it-1}$ , that is, the residuals of this regression,  $x_{it} = \sum_{j=1}^{k_i} \theta_{2ij} \Delta x_{it-j} + z_{it}\gamma_i + v_{it}$ . In the third step,  $\hat{e}_{it}$  is regressed on  $\hat{v}_{it-1}$ ,  $\hat{e}_{it} = \delta \hat{v}_{it-1} + \xi_{it}$ . The standard error,  $\hat{\sigma}_{ei}^2$ , of this regression is then used to normalize the residuals  $\hat{e}_{it}$  and  $\hat{v}_{it-1}$  (to control for heterogeneity in the variances of the series),  $\tilde{e}_{it} = \hat{e}_{it} / \hat{\sigma}_{ei}^2$ ,  $\tilde{v}_{it-1} = \hat{v}_{it-1} / \hat{\sigma}_{ei}^2$ . Finally,  $\rho$  is estimated from a regression of  $\tilde{e}_{it}$  on  $\tilde{v}_{it-1}$ ,  $\tilde{e}_{it} = \rho \tilde{v}_{it-1} + \xi_{it}$ . The conventional  $t$ -statistic for the autoregressive coefficient  $\rho$  has a standard normal limiting distribution if the underlying model does not include fixed effects and individual time trends ( $z_{it}$ ). Otherwise, this statistic has to be corrected using the first and second moments tabulated by Levin et al. (2002) and the ratio of the long-run variance to the short-run variance, which accounts for the nuisance parameters present in the specification. The limiting distribution of this corrected statistic is normal as  $N \rightarrow \infty$  and  $T \rightarrow \infty$ .

In contrast to the LLC test, the between-dimension-based panel unit-root test of Im, Pesaran, and Shin (2003) (IPS) allows the first-order autoregressive parameter to vary across countries by estimating the ADF equation separately for each country. Thus, the IPS test is less restrictive than the LLC test. The null hypothesis is that each series contains a unit-root,  $H_0 : \rho_i = 0$  for all  $i$ , and

the alternative hypothesis is that at least one of the individual series in the panel is (trend) stationary,  $H_1 : \rho_i < 0$  for at least one  $i$ .  $H_0$  is tested against  $H_1$  using the standardized  $t$ -bar test statistic

$$\Gamma_i = \frac{\sqrt{N}[\bar{t}_{NT} - \mu]}{\sqrt{\nu}}, \quad (\text{A.2})$$

where  $\bar{t}_{NT}$  is the average of the  $N$  cross-section ADF  $t$  statistics, and  $\mu$  and  $\nu$  are, respectively, the mean and variance of the average of the individual  $t$  statistics, tabulated by Im et al. (2003). The standardized  $t$ -bar statistic converges to a standard normal distribution as  $N$  and  $T \rightarrow \infty$ .

Both the LLC and the IPS test procedures assume cross-sectional independence and thus may lead to spurious inference if the errors,  $\varepsilon_{it}$ , are not independent across  $i$  (for example, due to common shocks or spillovers between countries). As discussed in Section 3.1, this issue might be of particular relevance for the total factor productivity series. Therefore, we also use the cross-sectionally augmented IPS test of Pesaran (2007), which allows for cross-sectional dependence by augmenting the ADF regression with the cross-section averages of lagged levels and first-differences of the individual series. The attractive feature of this test is that it permits the individual countries to respond differently to the common time effects as reflected by the country-specific coefficients on the cross-section averages of the variables. The cross-section augmented ADF (CADF) regression, carried out separately for each country, is given by

$$\Delta x_{it} = z_{it}\gamma_i + \rho_i x_{it-1} + \sum_{j=1}^{k_i} \phi_{ij} \Delta x_{it-j} + \alpha_i \bar{x}_{t-1} + \sum_{j=0}^{k_i} \eta_{ij} \Delta \bar{x}_{t-j} + v_{it}, \quad (\text{A.3})$$

where  $\bar{x}_t$  is the cross-section mean of  $x_{it}$ ,  $\bar{x}_t = N^{-1} \sum_{i=1}^N x_{it}$ . The cross-section augmented IPS statistic is a simple average of  $t_i$  defined by

$$CIPS = N^{-1} \sum_{i=1}^N t_i, \quad (\text{A.4})$$

where  $t_i$  is the OLS  $t$  ratio of  $\rho_i$  in the above CADF regression. Critical values are tabulated by Pesaran (2007).

Table A1 reports the results of these tests for the variables in levels and in first differences. As can be seen, all three test statistics are unable to reject the null hypothesis that  $\log(TFP_{it})$  and  $\log(OFDI_{it})$  have a unit-root in levels. Since the unit-root hypothesis can be rejected for the first differences, it can be concluded that all series are integrated of the same order (1) (that is,  $I(1)$ ), the necessary condition for cointegration in a bivariate context.

Table A.1

Panel unit root tests

Variables	Deterministic terms	LLC statistics	IPS statistics	CIPS statistics
Levels				
$\log(TFP_{it})$	$c, t$	10.52	0.15	-2.18
$\log(OFDI_{it})$	$c, t$	5.36	0.42	-2.02
First differences				
$\Delta\log(TFP_{it})$	$c$	-16.69**	-5.21**	-2.35**
$\Delta\log(OFDI_{it})$	$c$	-21.50**	-5.89**	-2.31*

Notes:  $c$  ( $t$ ) indicates that we allow for different intercepts (and time trends) for each country. Three lags were selected to adjust for autocorrelation. The relevant 1% (5%) critical value for the CIPS statistics is -2.83 (-2.67) with an intercept and a linear trend, and -2.32 (-2.15) with an intercept. \*\* (\*) denote significance at the 1% (5%) level.

## Appendix A2. Panel cointegration tests

We employ several techniques to tests for a cointegrating relationship between  $\log(TFP_{it})$  and  $\log(OFDI_{it})$ . First, we use the Pedroni (1999, 2004) framework which is based on a two-step estimation procedure. In the first step, the static cointegrating regression  $\log(TFP_{it}) = a_i + \delta_i t + b_i \log(OFDI_{it}) + \varepsilon_{it}$  is estimated separately for each country. Then, the estimated residuals,  $\hat{\varepsilon}_{it}$ , are tested for stationarity using seven test statistics. Four of these seven pool the autoregressive coefficients across different countries during the unit-root test and thus constrain the autoregressive parameters to be homogeneous across countries. Pedroni refers to these within-dimension-based statistics as *panel cointegration statistics*. The other three test statistics are based on estimators that average the individually estimated autoregressive coefficients for each country, thus allowing the autoregressive coefficient to be heterogeneous across countries. Pedroni refers to these between-dimension statistics as *group-mean panel cointegration statistics*. The first of the panel cointegration statistics is a non-parametric variance ratio test. The second and the third are panel versions of the Phillips and Perron (PP) *rho* statistic and *t* statistic, respectively. The fourth statistic is a panel ADF *t* test analogous to the LLC (2002) panel unit root test. Similarly, the first of the group-mean panel cointegration statistics is analogous to the PP *rho* statistic, the second is a panel version of the PP *t* statistic, and the third is a group mean ADF *t* test analogous to the IPS (2003) panel unit root test. The standardized distributions for the panel and group statistics are given by

$$\kappa = \frac{\varphi - \mu\sqrt{N}}{\sqrt{v}} \Rightarrow N(0, 1), \quad (\text{A.5})$$

where  $\varphi$  is the respective panel or group statistic, and  $\mu$  and  $\nu$  are the expected mean and variance of the corresponding statistic, tabulated by Pedroni (1999).

A potential problem with the Pedroni approach is that it does not take into account potential error cross-sectional dependence, which could bias the results (as discussed in Section 3.1). To test for cointegration in the presence of possible cross-sectional dependence we use the two-step procedure suggested by Holly et al. (2010). In the first step, we apply the common correlated effects (CCE) estimator of Pesaran (2006) to the static cointegrating regression. Like the cross-sectionally augmented IPS test, the CCE estimator allows for cross-sectional dependencies that potentially arise from multiple unobserved common factors and permits the individual responses to these factors to differ across countries. The cross-section augmented cointegrating regression for the  $i$ th cross-section is given by

$$\log(TFP_{it}) = a_i + \delta_i t + b_i \log(OFDI_{it}) + g_{1i} \overline{\log(TFP_t)} + g_{2i} \overline{\log(OFDI_t)} + e_{it}, \quad (\text{A.6})$$

where the cross-section averages  $\overline{\log(TFP_t)} = N^{-1} \sum_i \log(TFP_{it})$  and  $\overline{\log(OFDI_t)} = N^{-1} \sum_i \log(OFDI_{it})$  serve as proxies for the unobserved factors. In the second step, we compute the cross-section augmented IPS statistic for the residuals from the individual CCE long-run relations,  $\hat{\mu} = \log(TFP_{it}) - \hat{\delta}_i t - \hat{b}_i \log(OFDI_{it})$ , including an intercept. In doing so, we account for unobserved common factors that could be correlated with the observed regressors in both steps.

However, residual-based (panel) cointegration tests, such as the ones just described, restrict the long-run elasticities to be equal to the short-run elasticities. If this restriction is invalid, residual-based (panel) cointegration tests may suffer from low power (see, e.g., Westerlund, 2007). Another drawback of residual-based (panel) cointegration tests is that they are generally not invariant to the normalization of the cointegrating regression. Therefore, we also use the Larsson et al. (2001) procedure, which is based on Johansen's (1988) maximum likelihood estimation procedure. Like the Johansen time-series cointegration test, the Larsson et al. panel test treats all variables as potentially endogenous, thus avoiding the normalization problems inherent in residual-based cointegration tests. In addition, the Larsson et al. procedure allows the long-run elasticities to differ from the short-run elasticities and hence does not impose a possibly invalid common factor restriction. It involves estimating the Johansen vector error-correction model for each country separately:

$$\Delta y_{it} = \Pi_i y_{it-1} + \sum_{k=1}^{k_i} \Gamma_{ik} \Delta y_{it-k} + z_{it} \gamma_i + \varepsilon_{it}, \quad (\text{A.7})$$



where  $y_{it}$  is a  $p \times 1$  vector of endogenous variables ( $y_{it} = [\ln(TFP_{it}), \ln(OFDI_{it})]'$ ;  $p$  is the number of variables) and  $\Pi_i$  is the long-run matrix of order  $p \times p$ . If  $\Pi_i$  is of reduced rank,  $r_i < p$ , it is possible to let  $\Pi_i = \alpha_i \beta_i$ , where  $\beta_i$  is a  $p \times r_i$  matrix, the  $r_i$  columns of which represent the cointegrating vectors, and  $\alpha_i$  is a  $p \times r_i$  matrix whose  $p$  rows represent the error correction coefficients. The null hypothesis is that all of the  $N$  countries in the panel have a common cointegrating rank, i.e. at most  $r$  (possibly heterogeneous) cointegrating relationships among the  $p$  variables:  $H_0 : rank(\Pi_i) = r_i \leq r$  for all  $i=1, \dots, N$ , whereas the alternative hypothesis is that all the cross-sections have a higher rank:  $H_1 : rank(\Pi_i) = p$  for all  $i=1, \dots, N$ . To test  $H_0$  against  $H_1$ , a panel cointegration rank trace-test statistic is computed by calculating the average of the individual trace statistics,  $LR_{iT} \{H(r)|H(p)\}$ :

$$\overline{LR}_{NT} \{H(r)|H(p)\} = \frac{1}{N} \sum_{i=1}^N LR_{iT} \{H(r)|H(p)\}, \quad (\text{A.8})$$

and then standardizing it as follows:

$$\Psi_{LR} \{H(r)|H(p)\} = \frac{\sqrt{N} (\overline{LR}_{NT} \{H(r)|H(p)\} - E(Z_k))}{\sqrt{Var(Z_k)}} \Rightarrow N(0, 1). \quad (\text{A.9})$$

The mean  $E(Z_k)$  and variance  $Var(Z_k)$  of the asymptotic trace statistic are tabulated by Breitung (2005) for the model (with an intercept and a trend) we use. However, a problem is that the Johansen trace statistics tend to over-reject the null in small samples. To avoid the Larsson et al. test also overestimating the cointegrating rank, we compute the standardized panel trace statistics based on small-sample corrected country-specific trace statistics. Specifically, we use the small-sample correction factor suggested by Reinsel and Ahn (1992) to adjust the individual trace statistics as follows:

$$LR_{iT} \{H(r)|H(p)\} \times \left[ \frac{T - k_i \times p}{T} \right]. \quad (\text{A.10})$$

The results of these tests are presented in Table A2. Four of the seven Pedroni statistics reject the null of no cointegration at the 1% level. Specifically, the ADF-type tests reject the null hypothesis. Given that these tests have been shown to have the highest power for smaller sample sizes such as  $T = 26$  (see, e.g., Pedroni 2004), the ADF test results, in particular, provide strong evidence of cointegration. This conclusion is supported by the CIPS and the panel trace statistics which clearly show that  $\log(TFP_{it})$  and  $\log(OFDI_{it})$  are cointegrated (and exhibit a single cointegrating vector).

Table A.2

## Panel cointegration tests

Pedroni (1999, 2004) cointegration test statistics		
Panel $\nu$ statistic		1.05
Panel PP $\rho$ statistic		-1.76
Panel PP $t$ statistic		-8.66**
Panel ADF statistic		-6.37**
Group PP $\rho$ statistic		-0.17
Group PP $t$ statistic		-8.77**
Group ADF statistic		-5.28**
CIPS statistic for the residuals of the CCE long-run relations		
		-2.46**
Cointegration rank		
	$r = 0$	$r = 1$
Standardized panel trace statistics	6.04**	-0.27

Notes: \*\* indicate a rejection of the null of no cointegration at the one percent level. Under the alternative hypothesis, the panel  $\nu$  statistic diverges to positive infinity so that the right tail of the normal distribution is used to reject the null hypothesis. The relevant 1% critical value for the CIPS statistic is -2.32. The number of lags was determined by the Schwarz criterion with a maximum of three lags.

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