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Governments all over the world spend large amounts of resources in order to attract foreign direct investment (FDI), often based on the assumption that FDI increases total factor productivity (TFP) by bringing with it better technology and knowledge spillovers to domestic firms. However, evidence for this assumption from macro data is sparse. This paper is the first to examine the long-run effect of FDI on TFP using aggregate time-series data for the United States. The results show (i) that FDI is positively related to TFP in the long run, (ii) that FDI Granger-causes TFP growth in the long run, and (iii) that there is no long-run feedback from TFP to FDI.

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1. Introduction

Governments all over the world spend large amounts of resources in order to attract foreign direct investment (FDI), often based on the assumption that FDI increases overall, or total factor, productivity (TFP) by bringing with it better technology and knowledge spillovers to domestic firms. This assumption is articulated, for example, in the Monterrey Consensus of the UN Summit on Financing for Development, which states, inter alia, that "[FDI] is especially important for its potential to […] boost overall productivity […]" (United Nations, 2002: 9). However, evidence for this assumption from macro data is sparse and inconclusive.

While Woo (2009) finds in cross-sectional, pooled, and fixed effects regressions that FDI has an unconditional positive effect on TFP growth, the cross-country results of Alfaro *et al.* (2009) indicate that the impact of FDI on TFP growth is conditional on the level of financial development in the host country; the higher the level of financial development, the higher the effect of FDI on TFP growth. In contrast, Wang and Wong (2009), using panel seemingly unrelated regressions, find no robust evidence that the effect of FDI on TFP growth varies with the level of financial development. Their results suggest that the effect depends on the level of human capital in the host economy; more specifically, they find that FDI has a negative effect on TFP growth in countries with low levels of human capital, but the negative effect becomes smaller in absolute value and ultimately turns positive as the level of human capital increases. Because countries with low human capital tend to have low labor productivity and low absorptive capacity, this is inconsistent with the results of Baltabaev (2014), who, using panel GMM techniques, finds that FDI stimulates TFP growth only in countries where labor productivity relative to the United States is below a certain threshold. Finally, de Mello (1999) uses pooled, fixed-effects, and pooled mean-group regressions and reports mixed results: negative but mostly insignificant effects for his total sample of 15 OECD and 17 Non-OECD countries, positive effects for OECD countries, and statistically insignificant effects with different signs for Non-OECD countries.

One conclusion that can be drawn from these studies is that FDI can have positive, insignificant or even negative effects on TFP. Possible explanations for non-existent or negative effects (from the macro literature) include (among others) the following:[1] (i) To take advantage of knowledge spillovers from FDI, local firms need to reorganize their structure, buy new machines, and hire new managers and skilled labor. In countries with underdeveloped financial markets, domestic firms may be unable to make such investments (Alfaro *et al.*, 2009). (ii) The degree of productivity spillovers depends on the technology gap between foreign and local firms. On the one hand, it may be that indigenous firms need a certain level of absorptive capacity (in terms of human capital) before they can benefit from technologies developed by multinational firms. This would imply that mainly technologically advanced countries benefit from knowledge spillovers from foreign to domestic firms, while the role of FDI in technology transfer to domestic firms in less advanced countries is limited (Wooster and Diebel, 2010). On the other hand, it may be that the potential for positive spillovers is higher the larger the technology gap between foreign and local firms, which would imply that knowledge spillovers from FDI in technologically advanced countries are weak or non-existent (Wang and Blomström, 1992). (iii) Multinationals often source fewer inputs locally than the domestic firms they displace. This may lead to a decrease in local

[1] We refer here mainly to findings from multi-country *macroeconomic* studies on the FDI-TFP nexus, consistent with the macro focus of our analysis. These studies suggest (although not conclusively) that the effect of FDI on TFP may depend on the level of financial market development, the level of human capital, and the technology gap between foreign and local firms, as discussed above. However, it should be noted that the *microeconomic* literature on FDI spillovers suggests that the extent to which domestic firms benefit from these spillovers may also depend on several other factors, such as the level of trade openness, the level of property rights protection, and the mode of entry (joint ventures or wholly owned subsidiaries) (see, e.g., Iršová and Havránek, 2013).

demand for inputs and thus to a reduction in domestic input variety and hence lower productivity (Rodríguez-Clare, 1996). (iv) Multinationals have lower marginal costs due to some firm-specific advantage, which allows them to attract demand away from domestic firms, thus forcing the domestic firms to reduce production and move up their average cost curve (Görg and Greenaway, 2004). This may reduce not only the productivity of domestic firms but also the productivity of the economy as a whole (depending on relative productivity of foreign firms and the amount of the reduction of the productivity of domestic firms).

Another conclusion that can be drawn from the studies summarized above is that the effect of FDI on TFP is very likely to differ across countries. However, the evidence is too sparse and ambiguous to allow generalizations about the impact of FDI on TFP in specific countries. This motivates the present study, in which we examine the long-run effect of FDI on TFP using aggregate time-series data for the United States over the period 1980-2011.

The US analysis is potentially of international interest for several reasons. First, the United States is not only the most important source of FDI but also tops the list of host countries of foreign-based multinationals (Chintrakarn *et al.* 2012). Second, the United States is generally viewed as the most technologically advanced country of the world (Cantwell and Vertova, 2004). Finally, the United States has well-developed financial markets (Maskus *et al.*, 2012). Thus, the location of FDI in the United States provides a quasi-experiment to assess the general impact of FDI on TFP in countries with well-developed financial markets and small technology gaps between foreign investors and domestic firms, many of which are already operating close to the international technology frontier.

Although this paper presents the first macro study of the long-run effect of FDI on TFP in the United States,[2] there are related micro studies on FDI spillovers using US manufacturing data. Branstetter (2006) and Keller and Yeaple (2009), for example, find positive spillovers from multinationals to domestic firms. Mullen and Williams (2007), in contrast, find no evidence of positive spillovers; some of their findings suggest that FDI may actually reduce domestic firm productivity. However, because these studies focus on the manufacturing sector, and thus exclude the service sector, they do not capture potential spillovers from FDI in services, which accounts for the largest share of the total stock of inward FDI in the United States (about 60% in 2011 according OECD data);[3] nor do they capture potential FDI spillovers between the manufacturing and service sectors. Moreover, by their nature, such micro studies do not capture the direct positive effect of the higher productivity of foreign firms on the productivity of all firms. The relevant point in this context is that, given that foreign firms are generally more productive than domestic firms, the long-term net effect of FDI on aggregate productivity can be positive even if the foreign firms reduce the productivity of domestic firms. In this sense, the purpose of this macro study is to complement the existing micro studies of FDI spillovers in the United States, by examining the effect of FDI on TFP in the United States for the economy as a whole using cointegration and causality analysis. The study shows that FDI has a positive long-run effect on TFP in the United States and that long-run causality runs in only one direction, from FDI to TFP

The rest of this paper is organized as follows. Section 2 presents the basic empirical model and the data. The empirical analysis is presented in Section 3, and Section 4 concludes.

[2] It should be noted that de Mello (1999) estimates the effects of FDI on TFP growth using both panel and time-series data. While his panel analysis includes the United States, the United States is excluded from his multivariate time-series analysis of the effects of FDI on the growth rate of TFP; in this analysis, he considers only countries with stationary time-series data (Italy and Venezuela). Apart from this, he employs the vector error correction model approach of Johansen (1988) in his multivariate time-series analysis, an approach which is inappropriate and inaccurate when the variables are stationary.

[3] Available at https://stats.oecd.org/Index.aspx?DataSetCode=FDI_POSITION_INDUSTRY.

2. Model and data

The basic model is as follows:

$$\log(TFP_t) = c_i + \beta(FDI_t / j_t) + \varepsilon_t, \quad (1)$$

where $\log(TFP_t)$ is the log of TFP in period t and (FDI_t/j_t) represents two measures of FDI for period t. The first is the ratio of the stock of FDI to GDP, (FDI_t/GDP_t). Following most of the recent work on the effects of FDI (see, e.g., Ford *et al.*, 2008; de Sousa and Lochard, 2011; Chintrakarn *et al.*, 2012, Baltabaev, 2014), we use stocks rather than flows because stocks may more effectively capture long-run effects due to the accumulation of flows. More specifically, the use of FDI stocks ensures that the effects of FDI are not limited to the period in which the investment is made and thus that the effects of both new and established foreign firms are fully accounted for. Moreover, TFP is typically interpreted as a measure that captures (among other things) the stock of knowledge in an economy, implying that it is reasonable to assume a relationship between TFP and the stock, rather than the flow, of FDI. As is standard in the literature, FDI is scaled by GDP to control for market size effects, and thus to capture the extent of foreign presence in the economy. To ensure that our results are not due to external effects of capital accumulation per se, but to the higher productivity of foreign firms and their spillovers to domestic firms, the stock of FDI is also scaled by the total physical capital stock K_t. Thus, our second measure of FDI activity is given as (FDI_t/K_t). Both ratios are expressed in percentage terms. Because our dependent variable is in logs, this implies that $\beta\times100$ is the long-run semi-elasticity of TFP with respect to (FDI_t/GDP_t) and (FDI_t/K_t), respectively.

Following Hall and Jones (1999), we calculate (log) TFP as the residual from a Cobb-Douglas production function with capital and human capital-augmented labor:

$$\log(TFP_t) = \log(Y_t) - (1-\alpha_t)\log(K_t) - \alpha_t \log(L_t h_t), \quad (2)$$

where Y_t is output, K_t is capital input, $L_t h_t$ is human capital-augmented labor input, defined as the product of "raw" labor L_t and human capital per worker h_t, $(1-\alpha_t)$ is the capital share of income, and α_t is the labor share of income.

All data used to calculate TFP are from the Penn World Tables (PWT) version 8.1 (Feenstra *et al.*, 2015) (available at http://www.rug.nl/research/ggdc/data/pwt/). Y_t is measured by (real) GDP in constant 2005 dollars, K_t by the constant 2005 dollar value of the stock of (real) capital (constructed by the perpetual inventory method), L_t by total hours worked (annual hours worked per employed person times the number of employed persons), α_t by compensation of employees and self-employed relative to GDP, and h_t by

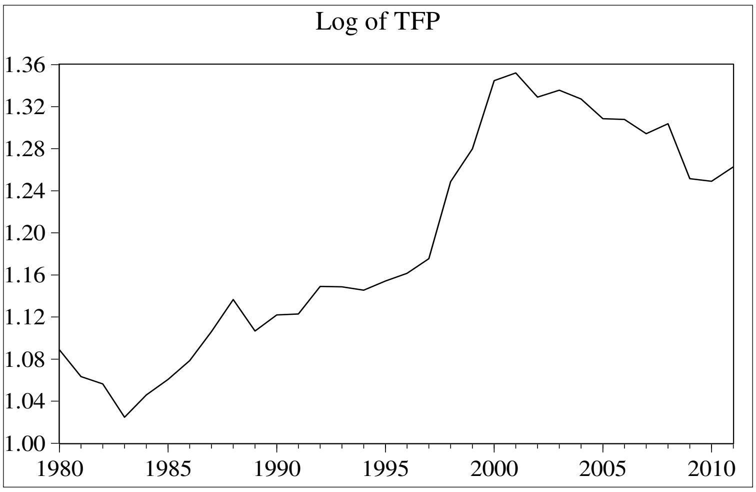
$$h_t = e^{\phi(s_t)}, \quad (3)$$

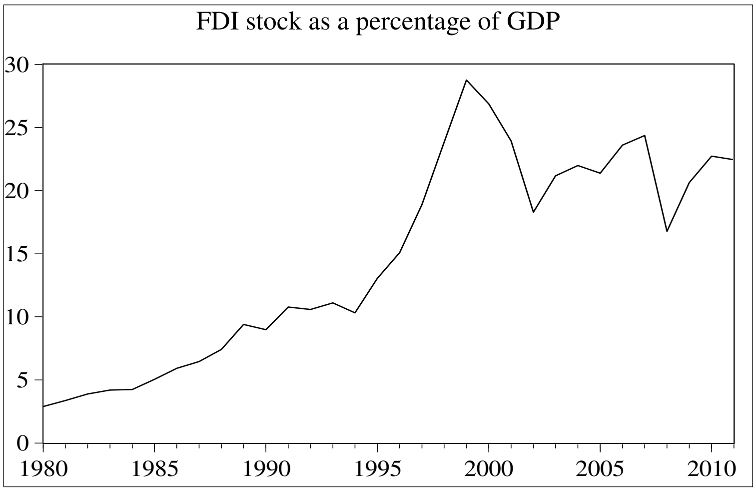
where s is the average years of schooling of the population above 15 years of age, the derivative $\phi'(s_t)$ is the return to schooling estimated in a Mincerian wage regression, and ϕ is a piecewise linear function (with a zero intercept and a slope of 0.134 through the fourth year of education, 0.101 for the next four years, and 0.068 for education beyond the eighth year).[4]

Data on the FDI stock/GDP ratio are from the UNCTADstat database (available at: http://unctadstat.unctad.org/wds/ReportFolders/reportFolders.aspx). FDI stock is 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 (UNTAD, 2008). To construct data on the ratio of the stock of FDI to the total capital stock, we divide the FDI stock/GDP ratio from the UNCTADstat database by the capital stock/GDP ratio from the PWT. Given that the UNCTAD data are available since 1980, while the PWT 8.1 data end in 2011, the empirical analysis covers the period from 1980 to 2011.

[4] The coefficient on the first four years is the return to schooling in sub-Saharan Africa (13.4%). The coefficient on the second four years is the world average return to schooling (10.1%). The coefficient on schooling above eight years is the OECD return to schooling (6.8%). All coefficients are taken from Psacharopoulos (1994).

Figure 1. TFP and FDI over the period 1980-2011





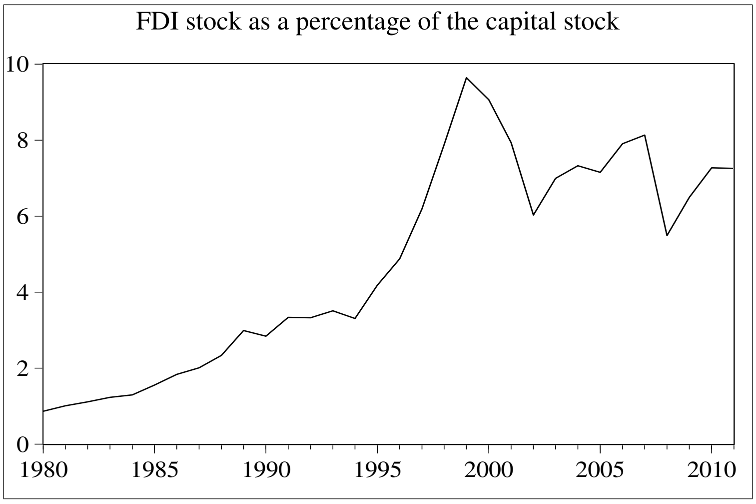


Figure 1 plots log(TFP_t), (FDI_t/GDP_t), and (FDI_t/K_t) for this period. As can be seen, all three variables are trending over time, implying that they are nonstationary. Given that most economic time series can be characterized by a stochastic trend model, it is likely that log(TFP_t), (FDI_t/GDP_t), and(FDI_t/K_t) also have stochastic, rather than deterministic, trends. If log(TFP_t) and (FDI_t/GDP_t), and log(TFP_t) and (FDI_t/K_t), respectively, are driven by two separate stochastic I(1) trends, then the linear combination of these nonstationary or integrated variables will also be I(1).[5] In this case, Equation (1) is a spurious regression in the sense of Granger and Newbold (1974), and there is no long-run relationship between TFP and FDI. If, however, log(TFP_t) and (FDI_t/GDP_t), and log(TFP_t) and (FDI_t/K_t), respectively, share a common stochastic trend, and no irrelevant nonstationary variables are included, then the linear combination of the variables will be stationary, or I(0). In this case, log(TFP_t) and (FDI_t/GDP_t), and log(TFP_t) and (FDI_t/K_t), respectively, are cointegrated in the sense of Engle and Granger (1987), and there exists a long-run relationship between permanent changes in TFP and permanent changes in FDI.

A well-known advantage of the cointegration framework is that, under cointegration, parameter estimates are superconsistent, and hence are robust to problems such as omitted variables and endogeneity (Coe *et al.*, 2009).

3. Empirical Analysis

3.1. Testing for cointegration

We use the autoregressive distributed lag (ARDL) bounds test methodology of Pesaran *et al.* (2001) to test for the existence of a long-run relationship between TFP and FDI. The advantage of this approach is that it applicable when it is not known with certainty whether the regressors are trend-stationary or integrated processes. To implement the bounds testing procedure, we specify the conditional error correction version of the ARDL model for Equation (1) as follows:

$$\begin{aligned}\Delta \log(TFP_t) &= b_1 + b_2 \log(TFP_{t-1}) + b_3\,(FDI_{t-1} / j_{t-1}) \\ &\quad + \sum_{i=1}^{k} \eta_i \Delta \log(TFP_{t-i}) + \sum_{i=0}^{k} \gamma_i \Delta(FDI_{t-i} / j_{t-i}) + u_t\end{aligned}, \tag{4}$$

where k is the lag length (which is determined by the Schwarz criterion). The null hypothesis of no cointegration is tested using both an F-test for the joint significance of the lagged levels of the variables (H_0: $b_2 = b_3 = 0$) and a t-test for the significance of the lagged level of the dependent variable (H_0: $b_2 = 0$). Pesaran *et al.* (2001) provide two sets of asymptotic critical values to test the null hypothesis: one when all variables are integrated of order one, I(1), and the other when the regressors are (trend) stationary, I(0). If the calculated test statistic lies above the upper bound critical value, then the null of no cointegration is rejected. If the test statistic is below the lower critical value, the null hypothesis of no cointegration cannot be rejected. If the test statistic falls between the lower and upper critical values, the result is inconclusive.

We use one lag of Δlog(TFP_t) and no lagged values of $\Delta(FDI_t/j_t)$, as suggested by the Schwarz Information Criterion, and also include an impulse dummy variable for 2003 to account for an outlier in the residuals and to ensure their normality. It should be explicitly noted that the results do not change qualitatively when the dummy is excluded; the dummy is included "only" to achieve normally distributed residuals.

Table 1 reports the calculated test statistics along with some residual diagnostics for both the error correction model of the relationship between log(TFP_t) and (FDI_t/GDP_t) and the error correction model of the relationship between log(TFP_t) and (FDI_t/K_t). *RESET* is the usual test for general nonlinearity and misspecification, *LM*(k), k = 1, 3 are Lagrange Multiplier (LM) tests for autocorrelation based on k lags, *ARCH*(k) is an LM test for autoregressive conditional heteroscedasticity, and *JB* is the Jarque-Bera test for normality. Since all p-values exceed the conventional significance levels, it can be concluded that

[5] The number in parenthesis denotes the order of integration. The order of integration is the number of times a time series must be differenced to make it stationary. Thus, an I(1) variable must be differenced once to make it stationary, or I(0).

neither obvious nonlinearity nor misspecification is present in both estimated models, and that the residuals of the two models show no signs of non-normality, autocorrelation or conditional heteroscedasticity. Thus, valid conclusions can be drawn from the results: both the *F*- and *t*-statistics are higher than the corresponding upper bound critical values for both models, implying that there is a (non-spurious) long-run relationship both between log(TFP_t) and (FDI_t/GDP_t) and between log(TFP_t) and (FDI_t/K_t).

Table 1. Bounds test for cointegration and diagnostic tests

Error correction model of the relationship between log(TFP_t) and (FDI_t/GDP_t)					
					Test statistics
F-statistic for the joint significance of the lagged level variables					24.98***
t-statistic for the significance of the lagged level of the dependent variable					-7.04***
Diagnostic tests					
RESET(1 (*F*-statistic)	*LM*(1) (*F*-statistic)	*LM*(3) (*F*-statistic)	*ARCH*(1) (*F*-statistic)	*ARCH*(3) (*F*-statistic)	*JB* ($\chi 2$-statistic)
0.75 (0.40)	1.29 (0.27)	0.79 (0.51)	0.0001 (0.98)	1.59 (0.22)	0.28 (0.89)
Error correction model of the relationship between log(TFP_t) and (FDI_t/K_t)					
					Test statistics
F-statistic for the joint significance of the lagged level variables					26.24***
t-statistic for the significance of the lagged level of the dependent variable					-7.21***
Diagnostic tests					
RESET(1 (*F*-statistic)	*LM*(1) (*F*-statistic)	*LM*(3) (*F*-statistic)	*ARCH*(1) (*F*-statistic)	*ARCH*(3) (*F*-statistic)	*JB* ($\chi 2$-statistic)
0.76 (0.46)	1.09 (0.31)	0.77 (0.52)	0.0003 (0.99)	1.48 (0.25)	0.57 (0.75)

1% critical value bounds		
	I(0)	*I*(1)
F-statistic	6.84	7.84
t-statistic	-3.43	-3.82

Notes: The critical value bounds are from Pesaran *et al.* (2001). The number in parentheses below the diagnostic test statistics are the corresponding *p*-values. One lag was used for the differenced TFP variable, and no lag for the differenced FDI variable. An impulse dummy for 2003 was included to achieve normally distributed residuals. *** indicate significance at the 1% level.

Table 2. ADF unit root and Johansen cointegration tests

ADF statistics		
	Levels (-4.28) [-3.56]	First differences (-3.67) [-2.96]
log(TFP_t)	-1.14	-4.34***
(FDI_t/GDP_t)	-1.80	-4.93***
(FDI_t/K_t)	-1.71	-4.63***
Johansen trace statistics		
	H_0: $r = 0$ (20.04) [15.41]	H_0: $r = 1$ (6.65) [3.76]
Test for cointegration between log(TFP_t) and(FDI_t/GDP_t)	20.95***	3.20
Test for cointegration between log(TFP_t) and(FDI_t/K_t)	21.89***	3.52

Notes: 1% critical values are in parentheses, and 5% critical values are in brackets. Critical values for the ADF tests are from MacKinnon (1991) and for the trace tests from Osterwald-Lenum (1992). The number of lags was determined using the Schwarz criterion with a maximum of five lags. The unit root tests for the levels include a constant and a trend, and the tests for the differences include only a constant. *r* denotes the number of cointegrating vectors. The trace statistics were adjusted using the small sample correction factor proposed by Reinsel and Ahn (1992) because in small samples the Johansen procedure can tend to over reject the null hypothesis of no cointegration in favor of finding cointegration. An impulse dummy for 2003 was included in the Johansen estimation to achieve normally distributed residuals. *** indicate rejection of the null hypothesis of a unit root or $r = 0$ (no cointegration) at the 1% level.

To check the robustness of this conclusion, we use the standard augmented Dickey-Fuller (ADF) (1979) unit root and Johansen (1988) cointegration trace tests. As is well known, the Johansen procedure can tend to over reject the null hypothesis of no cointegration in favor of finding cointegration in small samples (like the present one). Therefore, following standard practice, we adjust the trace test using the small sample correction factor proposed by Reinsel and Ahn (1992). The results of both tests are presented in Table 2. They show that the variables are integrated (of order one) and that log(TFP_t) is cointegrated with both (FDI_t/GDP_t) and (FDI_t/K_t).

3.2. Estimating the long-run relationship

The long-run parameter β in Equation (1) can be obtained from the error correction model (ECM) given by Equation (4) by dividing the estimated coefficient on the lagged level of the independent variable by the estimated coefficient on the lagged level of the dependent variable. The first column of Table 3 shows the resulting coefficients for both FDI variables. They are positive and significant at the 1% level, suggesting FDI has a positive long-run effect on TFP. The point estimates imply, if viewed causally, that, in the long run, a one percentage point increase in the FDI-to-GDP ratio and the FDI-to-capital-stock ratio, respectively, increases TFP by 1.2 percent and 3.7 percent, respectively (holding all else constant). Thus, our estimates are not only statistically but also economically significant.

To test the robustness of our estimates, we re-estimate the long-run relationship both between log(TFP_t) and (FDI_t/GDP_t) and between log(TFP_t) and (FDI_t/K_t) using three alternative estimation strategies: the DOLS method of Stock and Watson (1993), the FMOLS estimator of Phillips and Hansen (1990), and the maximum likelihood (ML) approach of Johansen (1988). As can be seen from Table 3, the four methods produce almost identical estimates of the effect of FDI on TFP for each measure of FDI.

Table 3. Estimates of the long-run effect of FDI on TFP

	ECM	DOLS	FMOLS	ML
Coefficient on (FDI_t/GDP_t)	0.012***	0.013***	0.013***	0.012***
	(17.82)	(17.59)	(13.51)	(17.14)
Coefficient on (FDI_t/K_t)	0.037***	0.038***	0.037***	0.037***
	(18.66)	(18.81)	(14.84)	(18.18)

Notes: The DOLS regression was estimated with one lead and one lag. One lag (chosen by the Schwarz criterion) was used in the Johansen procedure. *t*-statistics are in parenthesis. The *t*-statistic on the long-run coefficient of (FDI_t/j_t) from the error correction model was calculated using the delta method. The ECM and ML estimations include an impulse dummy for 2003. *** indicate significance at the 1% level.

3.3. Testing for long-run causality

The above interpretation of the estimation results is based on the assumption that long-run causality runs from (FDI_t/j_t) to log(TFP_t). However, given that multinationals often tend to be attracted to countries that have higher productivity, it is possible that that causality also runs in the opposite direction. In order to examine the direction of causality between FDI and TFP over time, we use the residuals (lagged one period) from the cointegrating equations for the two measures of FDI, $ec_t = \log(TFP_t) - 0.012(FDI_t / GDP_t)$ and $ec_t = \log(TFP_t) - 0.037(FDI_t / K_t)$, as error-correction terms in a vector error correction model (VECM) of the form

$$\begin{bmatrix} \Delta \log(TFP_t) \\ \Delta(FDI_t / j_t) \end{bmatrix} = \begin{bmatrix} c_1 \\ c_2 \end{bmatrix} + \sum_{i=1}^{k=3} \Gamma_i \begin{bmatrix} \Delta \log(TFP_{t-i}) \\ \Delta(FDI_t / j_t) \end{bmatrix} + \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} ec_{it-1} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}. \quad (5)$$

The Granger Representation Theorem (Engle and Granger, 1987) implies that for a long-run equilibrium relationship to exist between log(TFP_t) and (FDI_t/j_t) at least one of the α-

coefficients must be nonzero. If the adjustment coefficient α_1 in the $\Delta\log(TFP_t)$ equation is nonzero, then the null hypothesis of weak exogeneity is rejected for $\log(TFP_t)$. If the adjustment coefficient α_2 in the $\Delta(FDI_t/j_t)$ equation is nonzero, then the null hypothesis of weak exogeneity is rejected for (FDI_t/j_t). Hall and Milne (1994) show that weak exogeneity in a cointegrated system is equivalent to the notion of long-run noncausality. Thus, if (and only if) α_1 is nonzero, then (FDI_t/j_t) has a causal impact on $\log(TFP_t)$ in the long run; [6] if (and only if) α_2 is nonzero, then $\log(TFP_t)$ has a long-run causal impact on (FDI_t/j_t); if both α_1 and α_2 are nonzero, then long-run Granger causality runs in both directions.

Following common practice (see, e.g., Herzer, *et al.* 2012), we test for weak exogeneity and thus for long-run Granger non-causality by first eliminating the short-run dynamics in the model successively according to the lowest *t*-values and then using a conventional likelihood ratio test of the null hypothesis $\alpha_{1,2} = 0$.

Table 4 presents the results. The null hypothesis of weak exogeneity is rejected for $\log(TFP_t)$ at the 1% level both in the VECM with (FDI_t/GDP_t) and in the VECM with (FDI_t/K_t), suggesting that long-run causality runs from FDI to TFP. In contrast, the null hypothesis of weak exogeneity cannot be rejected for both FDI variables, indicating that there is no evidence of long-run causality from TFP to FDI.

Table 4. Tests for long-run causality

Test for long-run causality between $\log(TFP_t)$ and (FDI_t/GDP_t)		
	Weak exogeneity of $\log(TFP_t)$ (α_1)	Weak exogeneity of (FDI_t/GDP_t) (α_2)
$\chi^2(1)$	62.56	0.003
(p-values)	(0.000)	(0.96)
Test for long-run causality between $\log(TFP_t)$ and (FDI_t/K_t)		
	Weak exogeneity of $\log(TFP_t)$ (α_1)	Weak exogeneity of (FDI_t/K_t) (α_2)
$\chi^2(1)$	67.46	0.001
(p-values)	(0.000)	(0.98)

Notes: The number of degrees of freedom υ in the standard $\chi^2(\upsilon)$ tests corresponds to the number of zero restrictions. An impulse dummy for 2003 was included to achieve normally distributed residuals.

4. Conclusion

In this study, we examined the long-run relationship between total FDI and aggregate TFP in the United States over the period 1980-2011. Using cointegration and causality techniques, we found (i) that FDI is positively related to TFP in the long run, (ii) that FDI causes TFP growth in the long run, and (iii) that there is no long-run feedback from TFP to FDI. From these results, from the US case, we may cautiously conclude that countries with well-developed financial markets and small technology gaps between foreign investors and domestic firms tend to realize aggregate productivity gains from FDI. However, the exact extent to which these results can be generalized to other countries in similar circumstances is an open question.

Another open question is which factors determine the sign and magnitude of the effect of FDI on TFP in specific countries. Is it the level of financial market development? Is it the level of absorptive capacity? Or are there other, potentially more important, factors that play a role in determining the effect of FDI on TFP? For example, the productivity effect of FDI might depend to a large extent on the institutional, policy, and regulatory environment in

[6] This interpretation is based on the assumption that the cause occurs before the effect. Thus, this test cannot rule out the possibility that the (correct) expectation of the future level of TFP causes FDI in the long run. However, we consider this possibility unlikely because it is difficult, if not impossible, to predict the (correct) future level of TFP over long periods of time.

which firms operate. It can also be hypothesized that the effect of FDI on TFP depends on the level of trade openness. [7] Or it could be that FDI has a significant positive effect on TFP only if FDI exceeds a certain minimum level.

All these factors may induce apparent differences in the effect of FDI on TFP across countries. However, this single-country study is by its nature unable to determine which of these factors are important for explaining potential cross-country variations in the effectiveness of FDI.

Another inherent limitation of this study is that, due to its highly aggregated nature, it cannot provide any information as to which type of FDI (vertical or horizontal) has a larger impact on TFP and whether (and how) the effects of FDI on TFP differ across sectors. In this sense, and as mentioned in the Introduction, this study should be viewed as a complement, and not as an alternative, to studies that use micro data to explore the productivity effects of FDI.

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[7] The rationale underlying this hypothesis is that local firms in more open economies have more contact with firms from abroad than their counterparts in relatively closed economies, which allows them to acquire the necessary knowledge and skills to be able to learn from foreign investors. In addition, firms exposed to international competition are better able to compete with or supply multinationals (see, e.g., Iršová and Havránek, 2013).

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