A Dynamic Approach to the FDI-Environment Nexus:

The Case of China and India

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Abstract: The cointegration analysis and a vector error-correction (VEC) model are applied to examine the short- and long-run relationships among foreign direct investment (FDI), economic growth, and the environment in China and India. The results show that FDI inflow plays a pivotal role in determining the short- and long-run movement of economic growth through capital accumulation and technical spillovers in the two countries. However, FDI inflow in both countries is found to have a detrimental effect on environmental quality in both the short- and long-run, supporting *pollution haven hypothesis*. Finally, it is found that, in the short-run, there exists a unidirectional causality from FDI inflow to economic growth and the environment in China and India ─ a change in FDI inflow causes a consequence change in environmental quality and economic growth, but the reverse does not hold.

Key words: China, cointegration analysis, environment, FDI, India, vector error-correction

INTRODUCTION

Since the economic reform and opening up to the outside world in the late 1970s and the early 1980s, China and India have been the fastest growing economies in the world. Between 1992 and 2005, for example, the Chinese and Indian economies have grown on average by approximately 10% and 7% annually (Figure 1). Accordingly, foreign direct investment (FDI) inflows to the two countries have grown rapidly during the same period. Between 2000 and 2005, for example, the average annual inflows of FDI in China and India have reached $54.5 billion and $5.2 billion, respectively, more than double the amount of the 1992-1999 period (Table 1). As a result, China and India have become the largest and the ninth-largest FDI recipients (in terms of annual FDI inflows) among developing countries during the period of 2000-2005. However, the unprecedented economic growth in both countries over the past 25 years has been accompanied by obvious environmental pollution problems. Between 1980 and 2000, for example, sulfur dioxide (SO_2) emissions in China and India have increased by approximately 50% and 110%, respectively (Figure 1). Although some improvement achieved in the late 1990s due to the reinforcement of pollution control policies, China has become the largest emitter of SO_2 in the world. The primary objective of this paper is to examine the possible relationship between the inflow of FDI and environmental quality in China and India.

A plethora of studies has been conducted to deal with the economics of FDI in developing countries over the last three decades. Theoretical research in this area can be roughly categorized into two groups. The first group of studies has provided the theoretical rationale of the effect of FDI inflows on economic growth, which is known as

the FDI-growth nexus (e.g., Romer 1986, Lucas 1988, Rebelo 1991, Helphman and Grossman 1991). For example, the modern endogenous growth theory shows that long-run economic growth of the economy can result from more open liberalized government policies conductive for FDI inflows. More specifically, if capital is considered as knowledge rather than just plant and equipment, then the inflow of foreign capital can itself result in technological change and spillovers of ideas across countries (Grossman and Helphman 1991). With the capital exhibiting such increasing returns to scale, therefore, changes in FDI inflows can be an important vehicle for long-run economic growth in developing countries. The second group of studies has attempted to relate theoretical consideration to the impact of FDI on the environment in developing countries, which is referred to as the FDI-environment nexus (e.g., Pethig 1976, Copeland and Taylor 1994 and 1995, Porter and van der Linde 1995). For example, the pollution haven model asserts that, under globalization circumstance, the relatively lax environmental standards in developing countries become attractive comparative advantage to the pollution-intensive foreign capital seeking for weaker regulations to avoid paying costly pollution control compliance expenditure domestically (Copeland and Taylor 2003). On the other hand, the Porter hypothesis claims that, since environmental quality is a normal good, as income increases with FDI inflows, developing countries tend to adopt more strict environmental regulations (Porter and van der Linde 1995).

To date, on the other hand, empirical studies have mostly concentrated on how the inflow of FDI affects economic growth in developing countries (e.g., Tsai 1991, Wang and Swain 1997, Liu et al. 1997, Sun and Parikh 2001, Bende-Nabende et al. 2001, Liu et al. 2002, Shan 2002, Chakraborty and Basu 2002, Yao 2006, and Chang 2007). For

example, Wang and Swain (1997) employ a single equation model (i.e., ordinary least squares) to analyze factors affecting foreign capital inflows into China and Hungary; they show a positive relation between changes in the level of GDP and the inflow of FDI in those countries. Sun and Parikh (2001) use a structural model (i.e., three least squares) to examine the relationship between inward FDI, exports and economic growth in China; they find that an increase in FDI (and exports) has a positive and significant impact on Chinese economic growth. Chakraborty and Basu (2002) adopt a non-structural time series model (i.e., vector error-correction) to explore the dynamic interaction between FDI and economic growth in India; they discover evidence that GDP has a significant positive effect on inflows of FDI for the Indian economy in both short- and long-run.

Accordingly, empirical analyses of the FDI-environment nexus in developing countries have received little attention. To the best of our knowledge, Smarzynska and Wei (2001), Xing and Kolstad (2002), Eskeland and Harrison (2003), and He (2006) are the only four empirical studies that have attempted to address this issue. For example, Xing and Kolstad (2002) examine the effect of the U.S. FDI on environmental quality in both developed and developing countries; they find that developing countries tend to utilize lenient environmental regulations as a strategy to attract dirty industries from developed countries. He (2006) explores the relationship between FDI and the environment in China; he discovers evidence that an increase in FDI inflow results in deterioration of environmental quality. However, these studies implicitly assume a one-way causality from measures of environmental quality/regulations (SO_2 and CO_2 emissions or pollution abatement cost) and/or economic growth (GDP) to FDI and adopt a structural model (i.e., reduced-form equations) to estimate the impacts of FDI based on

such causality. As such, previous studies have neglected the endogenous nature as well as the possible causal relationships between FDI (and economic growth) and environmental quality in a multivariate framework; that is, whether an increase in FDI in developing countries caused by their weaker regulations deteriorates environmental quality or, alternatively FDI related spillover of knowledge tends to improve environmental quality via economic growth. In other words, no study has dealt with dynamic movements of FDI (and economic growth) and environmental quality.[1]

The contribution of this study, therefore, is to examine the FDI inflow-environment nexus in a dynamic framework of multivariate time-series. For this purpose, we assess the short- and long-run relationships among FDI, sulfur dioxide (SO_2) emissions and GDP in China and India using the Johansen cointegration analysis and vector error-correction (VEC) model. The Johansen approach features multivariate autoregression and maximum likelihood estimation; this method is well suited to address the issue of endogeneity and causal mechanisms when variables used in the model are non-stationary and cointegrated. In addition, the cointegration test is used to find the long-run equilibrium relationships among the selected variables. Finally, the VEC model provides information on the short-run dynamic adjustment to changes in the variables within the model. This analysis will shed new light on dynamic interrelationships between FDI inflows, economic growth and the environment, and contribute to the empirical literature on FDI-environment nexus.

In the next section, the theoretical and empirical modeling of FDI-environment nexus is presented. This is followed by a description of data used in the analysis and a

discussion of unit root tests. The empirical results are discussed followed by some conclusions.

MODELING OF FDI-ENVIRONMENT NEXUS

Theoretical Framework

In examining the dynamic relationship between FDI, GDP and SO_2 emissions in China and India, we rely on a FDI-environmental policy model developed by Xing and Kolstad (2002). More specifically, in its simplest form the foreign direct investment (FDI) in the host country can be specified as follows:

$$FDI = f_1(GDP, Z_1, R^*) \tag{1}$$

where GDP is the gross domestic product of the host country, which is used as a proxy for the strength of the economy; Z_1 is a vector of exogenous variables affecting FDI inflows such as cost structures (i.e., labor costs) and differentials in rewards of factor services; and R^* is the environmental regulatory laxity. The relationship between GDP and FDI is expected to be positive, implying that economic growth is the most important determinant for FDI inflow to the host country ($\partial FDI / \partial GDP > 0$). The positive relationship between FDI and R^* indicates that lax environmental policy is more attractive to pollution-intensive FDI, thereby increasing polluting industries in the host country.

Similarly, the pollution (E) such as SO_2 emissions in the host country can be specified as follows:

$$E = f_2(GDP, Z_2, R^*) \tag{2}$$

where Z_2 is a vector of exogenous variables affecting pollution levels such as energy consumption and prices. In general, the relationship between GDP and pollution emissions is expected to be positive, indicating that an increase in the scale of economic activity through income growth necessarily brings about a proportionate increase in pollution *(*$\partial E / \partial GDP > 0$*)*. Defining environmental quality as a normal good, however, it is further hypothesized that pollution emissions decrease as rising income passes beyond a threshold level ($\partial E / \partial GDP < 0$). Economists call this relationship as the *Environmental Kuznets Curve* (*EKC*) (Grossman and Krueger 1991 and 1993). The relationship between pollution and R^* is expected to be positive, implying that lenient environmental regulations result in an increase in environmental degradation.

Assuming that f_2 is invertible in R^{*2}, equation (2) can be solved for R^* as a function of the other variables as follows:

$$R^* = f_3(GDP, Z_2, E) \tag{3}$$

Finally, we substitute equation (3) into equation (1), which yields the following relationship:

$$FDI = g(GDP, E, Z_1, Z_2) \tag{4}$$

The estimation of equation (4) is the basic approach of this study. It should be emphasized that the relationship between FDI and pollution emissions (or environmental regulatory laxity) in developing countries is ambiguous and uncertain. More specifically, if pollution-intensive foreign capitals move to developing countries with weaker regulations, then the inflow of FDI deteriorates environmental quality ($\partial FDI / \partial E > 0$). On the other hand, if developing countries rely on technology transfer through FDI from developed countries as a primary means of technology acquisition, the inflow of FDI

tends to enforce environmental regulations via economic growth, thereby improving environmental quality ($\partial FDI / \partial E < 0$).

Specification of Time-Series Models

To estimate the long-run relationship among FDI, GDP and SO_2 emissions, we use the maximum likelihood estimation procedure developed by Johansen (1988) and Johansen and Juselius (1992). More specifically, given a vector Y_t of n potentially endogenous variables, it is possible to model Y_t as the cointegrated vector autoregression (VAR) having up to k lags as follows:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + ... + \Gamma_{k-1} \Delta Y_{t-k+1} + \Pi Y_{t-k} + \mu + u_t \quad (5)$$

where Y_t is a (3×1) vector of endogenous variables, $Y_t = [SO2_t, GDP_t, FDI_t]$; Δ is the difference operator; $\Gamma_1, ..., \Gamma_{k-1}$ are the coefficient matrices of short-term dynamics; $\Pi = -(I - \Pi_1 + ... + \Pi_k)$ are the matrix of long-run coefficients, where I is the identity matrix; μ is a vector of constant; and u_t is a vector of normally and independently distributed error terms, or white noise. If the coefficient matrix Π has reduced rank — i.e., there are $r \leq (n-1)$ cointegration vectors present, then the Π can be decomposed into a matrix of loading vectors, α, and a matrix of cointegrating vectors, β, such as $\Pi = \alpha\beta'$, where r is the number of cointegrating relations, α represents the speed of adjustment to equilibrium, and β' is a matrix of long-run coefficients. For three endogenous non-stationary variables in our analysis, for example, the term $\beta' Y_{t-k}$ in equation (5) represents up to two linearly independent cointegrating relations in the system. The

number of cointegration vectors, the rank of Π, in the model is determined by the likelihood ratio test (Johansen 1988).

If all variable in a vector of stochastic process Y_t are cointegrated, an error-correction representation captures the short-run dynamics while restricting the long-run behavior of variables to converge to their cointegrating relationships (Engle and Granger 1987). This can be done by estimating an error-correction model in which residuals from the equilibrium cointegrating regression are used as an error-correcting regressor. For this purpose, equation (5) can be reformulated as a short-run dynamic model as follows:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{k-1} \Delta Y_{t-k+1} + \alpha(\beta' Y_{t-1}) + \mu + \varepsilon_t \qquad (6)$$

where $\beta' Y_{t-1}$ is a measure of the error or deviation from the equilibrium, which is obtained from lagged residuals from the cointegrating vectors. Since the series are cointegrated, equation (6) incorporates both short- and long-run effects. That is, if the long-run equilibrium holds, then the term $\beta' Y_{t-1}$ is equal to zero. During periods of disequilibrium, on the other hand, this term is non-zero and measures the distance of the system from equilibrium during time t. Thus, an estimate of α provides information on the speed-of-adjustment, which implies how the variable Y_t changes in response to disequilibrium.

DATA AND TESTING FOR UNIT ROOTSs

Data

It is worth noting that among principal air pollutants, sulfur dioxide (SO_2) and carbon dioxide (CO_2) are the major measures of air pollution that have been widely used in the empirical studies. Of those, SO_2 represents the measure of local air pollution, whereas

CO_2 represents a global pollutant (externality), which individual countries are unable to regulate without international cooperation (Frankel and Rose 2005, He 2006). Given our individual country-specific approach, therefore, it is more appropriate to select SO_2 emissions as a proxy for the measure of environmental quality in China and India.

Annual time-series data on sulfur emission (SO_2), GDP and inward foreign direct investment are collected for China over the period 1980-2002 and for India over the period 1978-2000, respectively. The estimated sulfur emissions (measured in thousand tons) for China and India are obtained from a large database constructed by David Stern (Stern 2005 and 2006), which is known as the David Stern's Datasite (available at the web site http://www.rpi.edu/~sternd/datasite.html). Note that the data on sulfur emissions (SO_2) used in empirical studies have almost invariably come from the ASL and Associates database (ASL and Associate 1997, Lefohn et al. 1999), which compiles annual time-series data on SO_2 emissions for individual countries from 1850 to 1990. However, the unavailability of data after 1990 has limited continued use of these estimates for further research. Hence, David Stern has developed global and individual country estimates of sulfur emissions from 1991 to 2000 or 2002 (most OECD countries, including China) combined with estimates from existing published and reported sources for 1850-1990 (see Stern (2005) for more details). The real GDP of China and India is measured as the real GDP index (2000=100) and is taken from the International Financial Statistics (IFS) Online Service provided by the International Monetary Fund (IMF) (available at the web site http://www.imfstatistics.org/imf/). The values of FDI for China and India are measured as the inward FDI flows (measured in million U.S. dollars) and are obtained from the World Investment Report (WIR) provided by the United Nations

Conference on Trade and Development (UNCTAD) GlobStat Database (available at the web site http://stats.unctad.org/FDI/ReportFolders/reportFolders.aspx). The inward FDI flows are deflated using the GDP deflators (2000=100) obtained from the IFS.

Testing For Unit Roots

When dealing with time-series data, the possibility of unit roots in a series raises issues about parameter inference and spurious regression (Wooldridge 2000). For example, OLS regression involving non-stationary series no longer provides the valid interpretations of the standard statistics such as t -statistics, F -statistics, and confidence intervals. To avoid this problem, non-stationary variables should be differentiated to make them stationary. However, Engle and Granger (1987) show that, even in the case that all the variables in a model are non-stationary, it is possible for a linear combination of integrated variables to be stationary. In this case, the variables are said to be cointegrated and the problem of spurious regression does not arise. As a result, the first requirement for cointegration analysis is that the selected variables must be non-stationary.

To determine the existence of a unit root in the series, we examine the integration order of individual time-series ($SO2_t, GDP_t, FDI_t$) for China and India using the Dickey-Fuller generalized least squares (DF-GLS) test (Elliot et al. 1996). This test optimizes the power of the conventional augmented Dickey-Fuller (ADF) test using a form of detrending. The DF-GLS test works well in small samples and has substantially improved power when an unknown mean or trend is present (Elliot et al. 1996). The results show that the levels of all the series are non-stationary, while the first differences are stationary (Table 2), indicating that the six variables are non-stationary and integrated of order one,

or $I(1)$. The DF-GLS test statistics are estimated from a model that includes a constant and a trend variable. The Schwert Criterion (SC) is used to determine lag lengths for the unit root test.

EMPIRICAL RESULTS

Johansen Cointegration Test

Before implementing the cointegration test, the important specification issue to be addressed is the determination of the lag length for the VAR model, because the Johansen procedure is quite sensitive to changes in lag structure (Maddala and Kim 1998). The lag length (k) of the VAR model is determined based on the likelihood ratio (LR) tests. This method compares the models of different lag lengths sequentially to see if there is a significant difference in results (Doornik and Hendry 1994). For example, the hypothesis that there is no significant difference between a one- and a two-lag model cannot be rejected for both China and India at the 5% significance level. Thus, one lag (k=1) is used for both countries in our cointegration analysis. Diagnostic tests on the residuals of each equation and corresponding vector test statistics support k =1 as the most appropriate lag length for the VAR model (Table 3).[3] More specifically, in the residual serial correlation and heteroskedasticity tests using the F -form of the Lagrange Multiplier (LM) procedure, the null hypotheses of no serial correlation and no heteroskedasticity cannot be rejected at the 5% significance level. In the residual normality test using the Doornik-Hansen method (Doornik and Hansen 1994), on the other hand, the null hypothesis of normality can be rejected for 3 individual series and the

system for China at the 5% significance level. However, non-normality of residuals does not bias the results of the cointegration estimation (Gonzalo 1994).

The Johansen cointegration procedure is applied to determine the number of cointegration relationships among the three variables. The results show that, for both China and India, the trace tests reject the null hypothesis of no cointegrating vector ($r=0$), but fail to reject the null hypothesis of one cointegrating vector ($r=1$) at the 5% significance level (Table 4).[4] The results suggest that, for both China and India, there exists a stable, long-run equilibrium relationship between FDI, GDP and SO_2 emissions. Note that the system specification tests based on F-tests indicate that a linear trend is necessary for the VAR model of China, but it is not necessary for the VAR model of India.

Having obtained one cointegrating vector in both China and India, we test long-run weak exogeneity to examine the possible causal relationships between FDI, GDP and SO_2 emissions. A weakly exogenous variable can be interpreted as a driving variable that pushes the other variables away from adjusting to long-run equilibrium, but is not influenced by the other variables in the model. This test is implemented by restricting a parameter in speed-of-adjustment to zero ($\alpha_i=0$). The results show that, for both China and India, the null hypothesis of weak exogeneity cannot be rejected for FDI at the 5% significance level, indicating that this variable is weakly exogenous to the long-run relationship in the model (Table 5). These findings suggest that, for China and India, FDI inflow is the driving variable in the system and significantly affect the long-run movement of economic growth, but are not influenced by economic growth. This further implies that FDI inflow plays a significant role in the long-run economic growth in the

two countries through capital accumulation and technical spillovers (transfers of knowledge and skills). On the other hand, SO_2 emissions are also found to be weakly exogenous at the 5% significance level in China, indicating that SO_2 emissions do not adjust to deviations from any equilibrium state defined by the cointegration relation. This further suggests that with relatively weaker environmental regulations, China tends to attract more capital inflow from developed countries for pollution intensive industries, which in turn leads to higher economic growth.

When determining the existence of cointegration relationship, the cointegration vectors (β_j) estimated from equation (5) represent the long-run relationship among the selected variables. More specifically, having obtained only one cointegration relationship between FDI, GDP and SO_2 emissions in both China and India, the first eigenvector (β_1) of the three eigenvectors is most highly correlated with the stationary part of the process ΔY_t when corrected for the lagged values of the differences. As such, β_1 represents the cointegration vector determined by the cointegrated VAR model (Johansen 1988). After normalizing the coefficients of FDI, for example, the long-run equilibrium relation (β_1) between the three variables in China and India can be represented as the reduced forms of equations (7) and (8), respectively:

$$FDI_t = 0.08GDP_t + 7.98SO2_t + 4.96trend \quad (7)$$

$$FDI_t = 0.21GDP_t + 2.04SO2_t \quad (8)$$

Equations (7) and (8) show that economic growth in China and India has a positive long-run relationship with FDI, indicating that economic growth tends to attract more FDI inflow. In addition, a positive long-run relationship between SO_2 emissions and FDI in both countries implies that an increase in SO_2 emissions (or relaxation of environmental

regulations) tend to an increase in FDI inflow. This finding provides supportive evidence for the so-called *pollution haven hypothesis*; that is, such developing countries as China and India tend to utilize lenient environmental regulations in an effort to attract multinational corporations, particularly those engaged in highly polluting activities from developed countries.

VEC Model

The VEC model is estimated to identify the short-run adjustment to long-run steady states as well as the short-run dynamics among FDI, GDP and SO_2 emissions in China and India. For this purpose, we estimate the short-run VAR model in equation (6), with the identified cointegration relationships in equations (7) and (8). We adopt a general-to-specific procedure to estimate the VEC model (Hendry 1995). In the case of China, for example, since FDI and SO_2 emissions are found to be weakly exogenous to the system, the VEC model is first estimated conditional on the two variables. By eliminating all the insignificant variables based on an F -test, the parsimonious VEC (PVEC) model is then estimated using OLS (Harris and Sollis 2003). Likewise, the VEC model for India is estimated conditional on FDI. The number of lags used in the PVEC model is the same as that in the cointegration analysis. The multivariate diagnostic tests on the estimated model as a system show no serious problems with serial correlation, heteroskedasticity, and normality (Table 6). This suggests that the PVEC specifications do not violate any of the standard assumptions.

The results show that the error-correction terms (EC_{t-1}) for China and India are negative and significant at the 5% significance level (Table 6). More specifically, the

negative coefficients of EC_{t-1} ensure that the long-run equilibrium can be achieved. The absolute value of EC_{t-1} indicates the speed of adjustment to equilibrium. As such, the results indicate that, with a shock to the Chinese and Indian economies, GDP and SO_2 emissions tend to recover to their long-run equilibrium position. However, the adjustment toward equilibrium is not instantaneous. For example, the coefficients of EC_{t-1} for the ΔGDP_t equations in China and India are -0.53 and -0.60, respectively, suggesting that approximately 53%-60% of the adjustment occurs in a year for both countries. On the other hand, the coefficient of EC_{t-1} for the $\Delta SO2_t$ equation in India is -0.03, indicating very slow rate of adjustment toward long-run equilibrium.

The coefficients of the lagged variables in the PVEC models show the short-run dynamics (causal linkage) of the dependent variables (Table 6). More specifically, ΔFDI_t and ΔFDI_{t-1} are significant for the ΔGDP_t equations in China and India, indicating that FDI inflow has a positive effect on economic growth through its influence on the changes in industrial technology. Additionally, for India, ΔFDI_t and ΔFDI_{t-1} are also significant for the $\Delta SO2_t$ equation, implying that FDI inflow, particularly of pollution intensive industries from developed countries, tends to increase SO_2 emissions in India. Overall, the short-run dynamics are characterized by unidirectional causation; that is, economic growth (and SO_2 emissions) significantly affected by the inflow of FDI to China (India), but the reverse does not hold.

CONCLUSIONS

In this paper, we examine both the short- and long-run relationships between FDI, economic growth and environmental quality in China and India. The main contribution of this paper is to directly deal with the issue of potential endogeneity problems and the possible causal mechanism of FDI, economic growth (measured by GDP) and environmental quality (measured by SO_2 emissions) in a multivariate time-series framework. For this purpose, we adopt the Johansen cointegration analysis and a VEC model.

The empirical results show a positive long-run relationship between FDI and GDP for both China and India; that is, FDI inflow tends to stimulate economic growth. We also find a positive long-run relationship between FDI and SO_2 emissions in both countries; that is, lax environmental policy tends to attract more FDI inflow of pollution intensive industries from developed countries. The results further show that FDI is weakly exogenous to the long-run relationship in the models for China and India; that is, FDI inflow plays a key role in determining the long-run movement of economic growth through its influence on technical change. Finally, in the short-run dynamics, only a unidirectional causal link exists running from FDI to GDP and SO_2 emissions in both China and India, implying that a change in the inflow of FDI causes a consequence change in the level of GDP and environmental quality, but the reverse does not hold.

Notes

1. Notably, He (2006) has directly addressed the endogeneity problem between FDI, economic growth and SO_2 emissions in his analysis. However, he also employs a structural econometric model (i.e., simultaneous systems of equations) based on panel dataset.
2. Since the environmental regulatory laxity is not directly observed, Xing and Kolstad (2002) solve this latent variable problem by using pollutant emissions to infer laxity. For example, SO_2 emissions can be used as a yardstick to characterize the change of environmental regulation laxity; that is, relaxation (enforcement) of environmental regulation leads to an increase (decrease) in SO_2 emissions. Accordingly, pollution emissions (E) and environmental regulatory laxity (R^*) is interchangeable in this model.
3. The sample size could be another issue of concern for the Johansen procedure, because finite-sample analyses can bias the cointegration test toward finding the long-run relationship either too often or too infrequently. In fact, the number of observations used in this study seems to be a bit small; our findings should thus be viewed with caution. However, Hakkio and Rush (1991) note: "Our Monte Carlo studies show that the power of a cointegration test depends more on the span of the data rather than on the number of observations. Furthermore, increasing the number of observations, particularly by using monthly or quarterly data, does not add any robustness to the results in tests of cointegration." Following these authors, the annual data used in this study (23 years) can be considered to be long enough to reflect the long-run relationship between FDI, GDP and SO_2 emissions, which should somewhat mitigate our concern with the relatively small sample size.

4. Doornik and Hendry (2001) note: “The sequence of trace tests leads to a consistent test procedure, but no such result is available for the maximum eigenvalue test. Therefore current practice is to only consider the former.” Following these authors, we only depend on the former to test the null hypothesis.

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Table 1. Inward foreign direct investment (FDI) in developing countries

	1992-1999 Average		*2000-2005 Average*	
	Inward FDI (Million $)	Share (%)	Inward FDI (Million $)	Share (%)
China	35,322	25.5	54,479	23.2
Hong Kong	10,754	7.8	29,443	12.5
Mexico	9,678	7.0	20,346	8.7
Brazil	12,141	8.8	19,197	8.2
Singapore	9,288	6.7	14,300	6.1
Russia	2,330	1.7	7,515	3.2
Korea	2,846	2.1	6,157	2.6
Thailand	3,400	2.5	5,300	2.3
India	1,857	1.3	5,242	2.2
Chile	3,971	2.9	5,008	2.1
Sub Total	91,586	66.2	166,987	71.0
Total	138,251	100.0	235,078	100.0

Source: World Investment Report, United Nations Conference on Trade and Development (UNCTAD) GlobStat Database.
Note: Total means the sum of inflow FDI in developing economies. Share indicates % shares of each country's inward FDI in developing economies total FDI.

Table 2. Results of Dickey-Fuller Generalized Least Squares (DF-GLS) unit root test

	Variable	*Level*	*First difference*	*Lag*	*Decision*
China	$SO2_t$	-2.248	-3.952**	1	$I(1)$
	GDP_t	-1.034	-4.146**	1	$I(1)$
	FDI_t	-2.295	-4.694**	1	$I(1)$
India	$SO2_t$	-1.739	-3.936**	2	$I(1)$
	GDP_t	-0.941	-3.731**	1	$I(1)$
	FDI_t	-1.538	-3.251**	1	$I(1)$

Note: ** and * indicate rejection of null hypothesis of non-stationarity at the 5% and 10% significance levels, respectively. The 5%, and 10% critical values for the DF-GLS, including a constant and a trend, are -3.190, and -2.890, respectively.

Table 3. Diagnostic tests for residuals from Johansen cointegration estimation

		Serial correlation	*Heteroskedasticity*	*Normality*
China	$\Delta SO2_t$	0.47	0.03	11.75
		[0.50]	[0.87]	[0.00]**
	ΔGDP_t	2.14	1.15	1.68
		[0.16]	[0.30]	[0.43]
	ΔFDI_t	0.13	0.18	8.79
		[0.72]	[0.67]	[0.01]**
	System	1.33	0.53	22.57
		[0.27]	[0.94]	[0.00]**
India	$\Delta SO2_t$	0.66	0.01	4.47
		[0.43]	[0.94]	[0.11]
	ΔGDP_t	0.26	0.17	7.37
		[0.62]	[0.69]	[0.03]**
	ΔFDI_t	1.37	0.32	2.50
		[0.27]	[0.58]	[0.29]
	System	0.80	0.86	10.39
		[0.62]	[0.66]	[0.11]

Note: Δ denotes the first differences of the variables. p-values are given in parentheses. ** and * indicate rejection of null hypothesis of non-stationarity at the 5% and 10% significance levels, respectively. Serial correlation of the residuals of individual equations and a whole system was examined using the F-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged independent variables. Heteroskedasticity was tested using the F-form of the LM test. Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hendry 1994).

Table 4. Results of Johansen cointegration rank tests

	Null hypothesis	*Eigenvalue*	*Trace statistics*
China	$H_0 : r = 0$	0.923	72.99 [0.00]**
	$H_0 : r = 1$	0.454	19.21 [0.28]
	$H_0 : r = 2$	0.266	6.51 [0.41]
India	$H_0 : r = 0$	0.757	42.99 [0.00]**
	$H_0 : r = 1$	0.365	13.33 [0.11]
	$H_0 : r = 2$	0.165	3.75 [0.12]

Note: ** indicates rejection of the null hypothesis at the 5% significance level. Parentheses are p-values. The trace test leads to a consistent test procedure, but the maximum eigenvalue test does not (Doornik and Hendry 2001, p. 175). For this reason, we only report the former to test the null hypotheses.

Table 5. Results of weak exogeneity tests

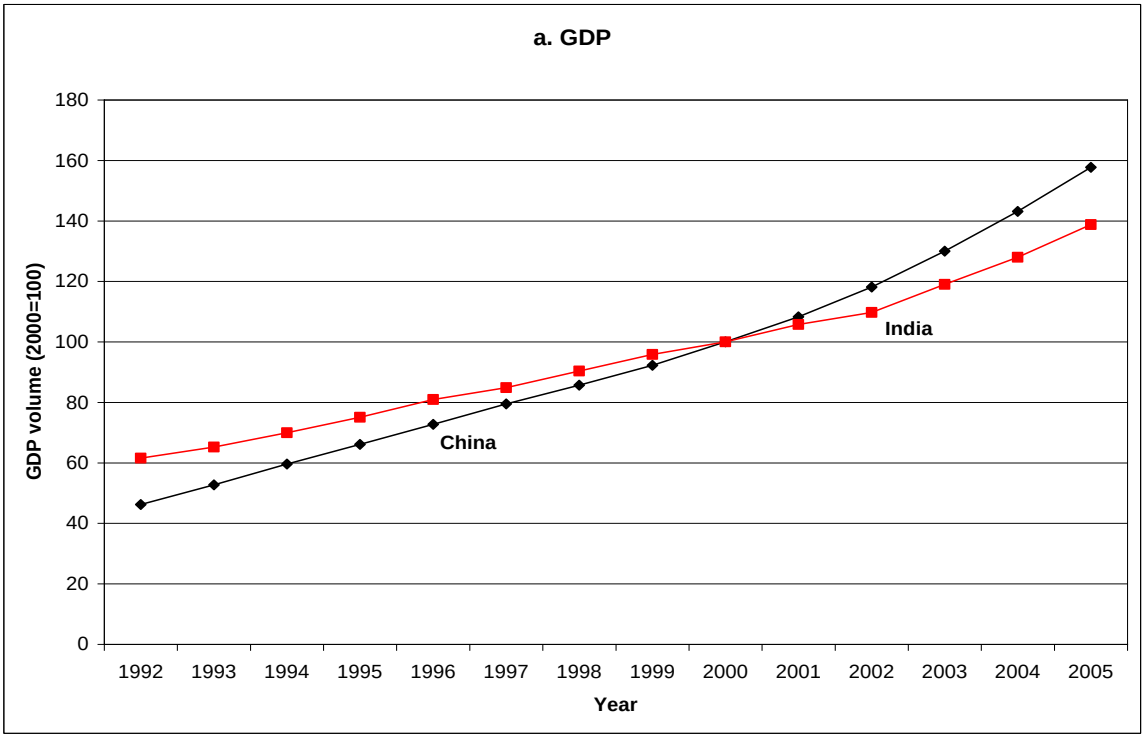
	Variable	*Weak exogeneity* $H_0 : \alpha_i = 0$
China	$SO2_t$	0.25 [0.62]
	GDP_t	36.34 [0.00]**
	FDI_t	0.10 [0.74]
India	$SO2_t$	6.06 [0.01]**
	GDP_t	14.74 [0.00]**
	FDI_t	0.23 [0.63]

Note: ** indicates the rejection of the null hypothesis at the 5% significance level. α_i represents the speed of adjustment to equilibrium. LR test statistic is based on the χ^2 distribution and parentheses are p-values.

Table 6. Results of parsimonious VEC models

	China	*India*	
	ΔGDP_t	$\Delta SO2_t$	ΔGDP_t
ΔFDI_t	0.12	0.08	0.03
	(3.65)**	(3.85)**	(0.09)
ΔFDI_{t-1}		0.07	0.67
		(2.90)**	(0.35)*
Constant	4.97	0.18	0.72
	(27.9)**	(5.60)**	(1.47)
EC_{t-1}	-0.53	-0.03	-0.60
	(-14.6)**	(-4.30)**	(-5.48)**
Serial correlation	1.77	0.66	
	[0.20]	[0.72]	
Heteroskedasticity	1.83	2.00	
	[0.20]	[0.31]	
Normality	0.66	7.22	
	[0.72]	[0.13]	

Note: ** and * indicate significance at the 5% and 10% levels, respectively. Parentheses in multivariate diagnostic tests are p-values.



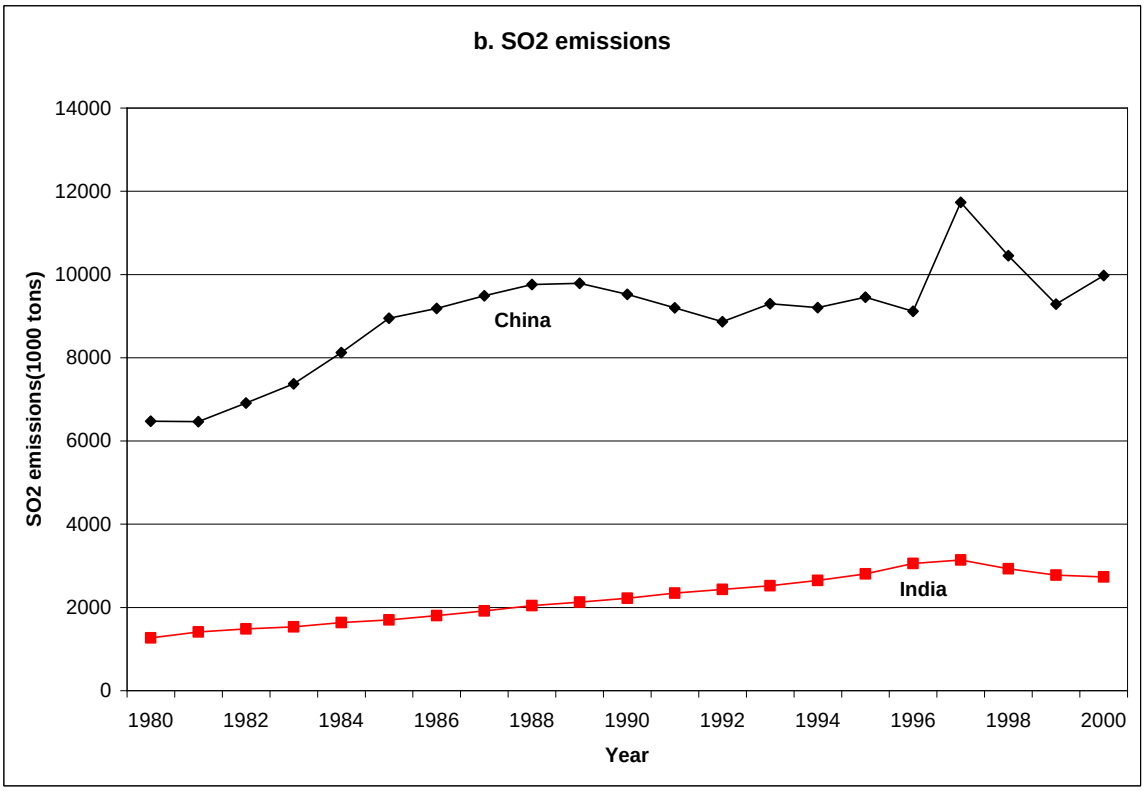


Figure 1. GDP and SO2 emissions in China and India