Quality of Institutions and Foreign Direct Investment in Developing Countries: Causality Tests for Cross-Country Panels

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June 4, 2006

Abstract

This paper analyzes the short-run and long-run dynamics between quality of institutions and foreign direct investment (FDI) in the sample of 62 developing countries covering the period 1984-2003. Panel cointegration test and FM OLS (Fully Modified OLS) estimators are used to test for cointegration. For short-run dynamics, we estimate error correction model using system GMM estimators. Institutional quality and FDI are found to have bi-directional cointegrating relationship in the long-run. However, there is no evidence in favor of short-run causality between two variables.

JEL *Classification:* C33, F21, F23, O17

Key Words: Quality of Institutions, Foreign Direct Investment, Cointegration, Short-run Causality; Developing Countries

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1. Introduction

The institutional quality of a host country received growing attention in the recent literature as one of the key determinants in location decisions made by foreign firms. Institutions provide the incentive structure for exchange that determines the cost of transaction and the cost of transformation in an economy (North, 1990). As an environment for investment, institutional variables such as the legal and political systems are thought to be crucial in affecting economic performance through their effect on investment decisions by curbing the risk of opportunism. This may be a particularly important issue for foreign investors who are not familiar with 'the rules of the game in a society' (Alfaro et al., 2005). Less corruption, and a fair, predictable, efficient bureaucracy may help attract FDI (Campos and Kinoshita, 2003).

The majority of papers on this topic provide evidence in support of a positive effect of the role of institutions in entry decisions by multinational enterprises (MNE). Campos and Kinoshita (2003) show that quality of institutions is one of the main determinants of FDI inflows to transition countries. Based on the analysis of panel data in developing countries, Gastanaga et al. (1998) demonstrate the direct effect of institutional characteristics on FDI. Using Japanese firm level data, Kang (2004) also finds that an institutional environment favorable to MNEs leads to a higher level of ownership of local companies. Wei (2000) examines a bilateral panel of FDI data and provides evidence that corruption in a host country negatively affects inward FDI particularly from U.S. and EU. Similarly, Aizenman and Spiegel (2000) introduce an imperfect enforcement contract framework and show that corruption discourages FDI more severely than it discourages domestic investments. Lee and Mansfield (1996) find that weakness of intellectual property protection in a host country can discourage FDI from the U.S. and influence more on the composition of U.S.

direct investment in wholly-owned foreign affiliates than partly-owned ones. Using data from 47 countries averaged over the sample period 1970-2000, Alfaro et al. (2005) find evidence that institutions are a robust casual determinant behind international capital mobility, partly offering an explanation of the 'Lucas paradox.'

On the other hand, little of the recent literature documents the possibility of a reverse link between institutional quality and FDI. Using geographical and cultural proximity to the major originators of FDI outflows as a set of instrumental variables, Larrain and Tavares (2004) provide evidence that FDI plays a significant role in lowering the level of corruption. One potential reason is that local officials and their behavior can be affected by foreign investors' standards of governance. Hellman et al. (2002) analyze the effect of FDI flows into transition economies. They suggest that FDI, instead of importing high standard of governance as it is believed, might provide a negative feedback to the host country by magnifying the problems of the state capturing and procuring kickbacks in a highly corrupt environment.

Nevertheless, the possibility of reverse causality was often neglected in the literature. Based on panel data measured in levels rather than differences, most related studies emphasize the unidirectional role of institutional quality in attracting FDI. However, a positive relationship between institutions and FDI does not automatically imply a true causal relationship. Critics can refer to the fact that there are a number of econometric problems with this approach because variables measured in the level form have common trends and measurement errors. Moreover, these statistical problems may be compounded by endogeneity problems if reverse causality is present (Zhang and Fan, 2001). The use of variables measured in differences, however, can also lead to the misspecification of the final model since it ignores any long-term relationship and captures only the short-term impact. Engle and Granger (1987) suggest a clever way to reconcile this problem by proposing the two-step estimation of error correction model (ECM). This approach is useful in investigating both long-run co-integration and a short-run Granger causality.

Causality is an important issue. Although both foreign technology (transferred primarily by FDI) and good institutions are regarded as the main two requirements of economic growth (Hausmann and Rodrik, 2003), the causal relationship between these variables has not been studied adequately. The study on this issue can facilitate explanation of the way these factors interact to foster economic growth, even though the growth is not the main topic in the present paper.

The purpose of this paper is to investigate both the long-run cointegrating relationship and short-run dynamics between FDI and the quality of institutions. Cointegration refers to a linear combination of non-stationary variables, implying that their stochastic trends must be linked as a long-run equilibrium. Short-run dynamics enable us to test the direction of short-run causality. This paper uses the bivariate error correction model (ECM) approach to apply a series of econometric tests to a panel of 62 developing countries observed over the period 1984-2003.

To the best of our knowledge, this is the first study that applies the ECM to examine the causality between FDI and institutional quality in developing countries. None of the former studies have explicitly examined the causality between FDI and the quality of institutions. In addition, most of the work is concentrated on the effect of corruption rather than institutions in a broader sense. Campos and Kinoshita (2003) used the GMM dynamic panel estimator to support the presence of positive effect of institutions on FDI, but they did not consider the integration and co-integration properties of the data. It is uncertain whether their results represent a true structural long-run equilibrium relationship or a spurious one (Christopoulos and Tsionas, 2004).

Our results from alternative panel estimation tests show that there is indeed a cointegrated long-run relationship between inward FDI and institutional environment of host country. In the short run, there is no explicit causation between two variables.

The remainder of our paper is organized as follows. The econometric model and empirical approach are introduced in section 2. Section 3 explains the data. Section 4 reports the empirical results, and section 5 concludes.

2. The Model and Econometric Technique

To investigate the causal relationship between FDI and institutional quality, we use the following autoregressive-distributed lag model (ADL):

$$y_{i,t} = \beta_0 + \beta_1 y_{i,t-1} + \beta_2 y_{i,t-2} + \beta_3 x_{i,t} + \beta_4 x_{i,t-1} + \beta_5 x_{t-2} + \psi_t + \varepsilon_i + u_{i,t} \tag{1}$$

where $i = 1, ...62, t = 1, ...20$

$y_{i,t}$ is either the log of inward FDI stock ($FDI_{i,t}$) or institutional quality ($Inst_{i,t}$) of country i in year i, ψ_t is the period-specific parameter capturing aggregate global shocks (assuming somewhat unrealistically that the sensitivities of $y_{i,t}$ to the shocks are identical for all countries), ε_i is an unobserved country-specific effect, and $u_{i,t}$ is a stochastic error term. The institutional quality index is measured as an average of the measures of three subcomponents of institutions-corruption, law and order, and bureaucracy quality. When the dependent variable is quality of institutions, ψ_t is dropped, since institutional variables are usually persistent and may not be significantly affected by worldwide macroeconomic shock. The lag length of two is chosen according to the AIC (Akaike Information Criteria) and BIC (Bayesian Information Criteria) model selection criteria.

2.1. Testing for Unit Root

Before we examine the existence of cointegration, we must verify that FDI and $Inst$ are integrated of order one, or $I(1)$ in levels. We use the panel unit root test of Im et. al, (1995) for both level and first-differenced variables. The use of panel-based tests is necessary because the power of country-by-country time-series unit root tests may be quite low given the sample size and time span of the data (Christopoulos and Tsionas, 2004). For comparison, we also apply the KPSS country-by-country stationarity tests (Kwiatkowski et al., 1992).

Im, Pesaran, and Shin (1995) propose a unit root test (IPS test) under the null that all series have unit roots against the alternative that some series are stationary. If the null cannot be rejected, it is likely that all series have a unit root. The test is based on the cross-country average of the individual augmented Dickey-Fuller (ADF) t−statistics. This is widely used for panel data unit root tests and many or most aggregate economic time series are such that the null hypothesis of a unit root cannot be rejected as a result of an IPS test. However, it is worth noting that the null hypothesis of a unit root is not rejected unless there is strong evidence against it. The common failure to reject a unit root may reflect the fact that either the original data are not very informative about having a unit root, or standard unit root tests are not very powerful against relevant alternatives (Kwiatkowski et al., 1992). As an effort to complement the unit root test in determining whether data are stationary or integrated, Kwiatkowski et al. (1992) developed a one-sided LM-test for the null hypothesis of trend stationarity against the alternative of difference stationarity (unit root) with deterministic trend. However, the drawback of both IPS and KPSS unit root tests is that their basic models are developed under the assumption of $T \to \infty$, which can mislead asymptotic results for panels where T is relatively small, as in our case. Hadri and Larsson (2005) propose the tests for panel data under the assumption of limit theory that T is fixed and the number of groups N is allowed to go infinity. This makes the test more applicable to panels with a large number of cross-sections relative to short period of time, and improves the finite sample properties. Hadri and Larsson (2005) construct the panel stationarity test statistic using two finite-sample moments of the corresponding KPSS statistics based on a single time series. They show that asymptotic distributions of the test statistics are normal and also provide the evidence of substantially higher power of the tests than for the tests based on a single time series. We follow their method and derive the panel statistic as an alternative to IPS panel unit root test and country-by-country KPSS stationarity test.

2.2. Testing for Cointegration

As a precondition for causality test, we need to check cointegrating properties of variables. Causality must run in at least one direction if two variables are cointegrated (Engle and Granger, 1987). A conventional test for cointegration is Johansen's procedure with country-by-country pure time series data. Since, however, the size and the power properties of Johansen's test with small sample sizes can be severely distorted (Christopoulos and Tsionas, 2004), a panel cointegration test developed by Pedroni (1999a) is conducted in this paper. Consider the following panel regression:

$$y_{i,t} = \alpha_i + \beta_i \chi_{i,t} + u_{i,t} \qquad for \;\; t = 1, \ldots, 20, \quad i = 1, \ldots, 62 \tag{2}$$

where α_i is the member-specific intercept and $\chi_{i,t} = \chi_{i,t-1} + \varepsilon_{i,t}$. Denote the series of estimated residual $e_{i,t}$ from regression (2), which follows the autoregressive process as:

$$e_{i,t} = \gamma_i e_{i,t-1} + v_{i,t}$$

y_t and χ_t are cointegrated of order I (1,1), if the series of estimated residual $e_{i,t}$ is stationary. Pedroni's between-dimension statistic is constructed with the null of no cointegration γ_i=1 against the alternative hypothesis γ_i<1 for all i. In contrast, the alternative hypothesis for within-group statistics is $\gamma_i = \gamma$. Thus, the between-group mean statistic allows for a more flexible alternative hypothesis. It allows for the short-run dynamics, fixed effects, and even the cointegrating vectors to differ across panel members under the alternative hypothesis of a single cointegrating vector (Pedroni, 1999a). He derives seven different statistics. For a small sample size, the test based on the group ADF-statistic is the most powerful, followed by the test based on the panel v-statistic. Thus, we will adopt group ADF-statistic as criteria of accepting or rejecting the null.

2.3. Estimating the Long-Run Relationship: Fully Modified OLS

Following the cointegration tests, we apply the fully modified OLS (FMOLS) method to estimate the long-run relationship. It is well known that despite its super-consistency, OLS estimation yields asymptotically biased results, because the nonstationary regressors are endogenously determined in the I(1) case (Christopoulos and Tsionas, 2004). On the other hand, FMOLS produces asymptotically unbiased estimators, as the statistic is constructed to make corrections for endogeneity and serial correlations to the OLS estimator $\beta_{i,t}$. Another advantage of this approach is that it allows us to pool the long-run information contained in the panel while permitting the short-run dynamics and fixed effects to be heterogeneous across different members of the panel (Pedroni, 1999b). The alternative method to obtain a reliable estimator for cointegration is to estimate the model using the GMM estimator proposed by Arellano and Bover (1995) after resolving the problem of endogeneity of regressors in equations for cointegration, or modifying the conventional two-step error correction model into a one-step ECM. This method is used as a robustness check for estimation of ECM using the OLS estimator.

2.4. Error Correction Model

Short-run dynamics are an important issue as well as a long-run cointegrating relationship. Having established the presence of a structural long-run relationship, we proceed to estimate both long- and short-run causalities between variables in a single error-correction equation. The test is made on the basis of the Engle-Granger two-step methodology (1987). Since the ADL model does not make a distinction between long- and short-run effects, the basic model represented by equation (1) is linearly transformed into the error correction model such as:

$$\Delta y_{i,t} = \beta_0 + (\beta_1 - 1)\Delta y_{i,t-1} + \beta_3 \Delta x_{i,t} + (\beta_3 + \beta_4)\Delta x_{i,t-1} + \lambda(y_{i,t-2} - \phi x_{i,t-2}) + \psi_t + \varepsilon_i + u_{i,t} \qquad (3)$$

The first three non-constant terms capture short-run dynamics while the error-correction term represents deviations from the long-run equilibrium. The error-correction term is obtained by saving residuals of separate estimation of the long-run equilibrium. The second step is to estimate equation (3). The parameter λ can be interpreted as the speed of adjustment since its coefficient represents the rate at which short-run dynamics of FDI (or $Inst$) converge to the long-run equilibrium relationship. Apparently, λ must be significantly different from zero if the variables are cointegrated.

However, although the two-step method is widely used in recent literature, we also choose the generalized one-step ECM as an alternative formulation of the Engle-Granger two-step procedure. The OLS estimate ϕ from the regression of equation (3) is a consistent estimate, but can be biased in a finite sample. The Monte Carlo results conducted by Banerjee et al. (1998) indicate that the finite sample bias can be substantial compared to generalized ECM. By imposing the restriction of homogeneity on the lagged dependent variable and the explanatory variable, the generalized ECM corresponding to (3) is transformed as follows:

$$\Delta y_{i,t} = \beta_0 + (\beta_1 - 1)\Delta y_{i,t-1} + \beta_3 \Delta x_{i,t} + (\beta_3 + \beta_4)\Delta x_{i,t-1} + \lambda(y_{i,t-2} - x_{i,t-2}) + \eta x_{i,t-2} + \psi_t + \varepsilon_i + u_{i,t} \quad (4)$$

It is proven that the estimated coefficients of the error correction terms are not affected by ϕ in the error correction term (Banerjee et al., 1990). This implies that it is sufficient to interpret the significance of λ as a deviation from long-run equilibrium, even though it is not the true long-run elasticity.

To control for endogeneity of nonstationary variables, we will implement the system GMM method using ECM proposed by Yasar et al. (2004) instead of OLS estimator. The general way to deal with dynamic panel data is to apply first-differenced GMM estimators using the levels of the series lagged two periods or more as instrumental variables. However, when the number of time series observations is small, as in our case, the first-differenced GMM may behave quite poorly because lagged levels of the variables are only weak instruments for subsequent first-differences (Bond et al., 2001). This problem may be alleviated by introducing the system GMM estimator suggested by Arellano and Bover (1995) and Blundell and Bond (1998). Under the additional assumption that first-differences are not correlated with country-specific effects, the basic idea of system GMM is to combine both equations in first-differences, taking the lagged level variables as instruments, with equations in levels with lagged first-differences as instruments. To illustrate, considering a simple AR(1) model:

$$y_{i,t} = \alpha y_{i,t-1} + \beta x_{i,t} + \eta_i + v_{i,t}$$

$$|\alpha| < 1 \quad \text{for} \quad i = 1,\ldots,N \quad \text{and} \quad t = 2,\ldots T \qquad (5)$$

where $x_{i,t}$ is correlated with η_i and endogenous so as to satisfy $E[x_{i,t} v_{i,s}] \neq 0$ for $i = 1,\ldots,N$ and $s \leq t$. Then two moments conditions for system GMM are:

$$E[x_{i,t-s} \Delta v_{i,t}] = 0 \quad \text{for} \quad t = 3,\ldots,T\,, \quad i = 1,\ldots,N \quad \text{and} \quad s \geq 2, \qquad (6)$$

$$E[\Delta x_{i,t-s} v_{i,t}] = 0 \quad \text{for} \quad t = 1,\ldots,T,\ \ i = 1,\ldots,N \qquad (7)$$

To establish the validity of instrumental variables, specification tests are conducted. The first specification test is the Sargan test, of which the null is that there is no correlation between instruments and errors. The failure to reject the null can be viewed as evidence in favor of using valid instruments. The null hypothesis of the second test is that the errors are not serially correlated in a first-differenced equation. By construction, the differenced error term may be first-order serially correlated even if the original error term is not (Carkovic and Levine, 2002). Thus, if the null of no serial correlation of AR(2) model cannot be rejected, it can be viewed as evidence supporting the validity of instruments used.

3. Data

The data used in this paper represents a balanced panel of 62 developing countries between 1984 and 2003. Due to the unavailability of sufficiently long time series data, most transition economies are not included. The FDI variable is per capita inward FDI stock in millions of U.S. dollars. The data source is the UNCTAD (United Nations

Conference on Trade and Development) statistical database.

The institutional quality data is provided by ICRG (International Country Risk Guide) researcher’s dataset. Three subcomponents of institutional quality include corruption, law and order, and bureaucracy quality. The assessment of these components is based on the subjective analysis of the available information. For measurement of corruption, potential and actual corruption in the form of nepotism, excessive patronage, and secret party funding are also considered, as well as financial corruption such as bribe; law and order are the assessment of strength and impartiality of the legal system and the popular observance of the law; bureaucracy quality stands for strength and expertise to govern without drastic changes in policy or interruptions in government services (ICRG, 2005). Since we look into the bivariate causality rather than treating each subcomponent as a different variable, we use the average of the normalized three components as a proxy for quality of institutions. The range of normalized point is from zero to one, where higher score implies better institutions. By the nature of this construction, these variables are bounded above and below by random numbers, which makes it impossible for the series to be nonstationary. Thus, we transform the index using inverse logit function to allow it to vary without limit.

4. Empirical Results

4.1. Time-Series Properties

In this section, we present our empirical results. IPS panel unit root tests without trend reported in Table 2 support the presence of a unit root in both FDI and institutional quality across countries, as well as stationarity of their first differences. As seen in the table, the inclusion of trend does not change the results for both FDI and quality of institutions.

For comparison purposes, country-by-country KPSS level stationarity and trend stationarity test results for FDI and quality of institutions are reported in Table 3. (Statistics are not reported.) The lag truncation parameter is selected as 2 according to AIC and BIC criteria. It is shown that the FDI variables of 50 countries appear to have unit roots without trend. Tests for trend stationarity, however, indicate that the null of trend stationarity cannot be rejected for 23 countries at 10% significance level.

The results are also mixed for quality of institutions. Institutional quality is level stationary for 21 countries. In 28 countries, quality of institutions is found to be trend stationary. Considering this sufficiently high number of countries with I(0) at individual test, it is not very safe to conclude that the results for both variables are consistent with the null of having unit root. However, since we have only 20 years of time span, unit root test results for individual time series are not very reliable. The number of countries with KPSS tests in first differences for both series is also reported in Table 3. Except eight countries, differenced FDI is stationary around level. Similarly, institutional quality in first difference is found to be level stationary for 52 countries. Figure 1-4 (histogram) and Appendix 2 (quantile of KPSS) represent the distribution of KPSS statistics of differenced variables with lag1 and lag2. There is no clear sign of the presence of a high level of heterogeneity across countries.

KPSS panel results are reported in Table 4. Since its assumption is fixed T and large N, it is more applicable for our data set. Moreover, the panel level test is more powerful than tests for individual countries. Both FDI and index of institutional quality turn out to have a unit root and their differenced series are stationary. Trend is not included to test for differenced series, as the trend term disappears when the original series is differenced. Thus, despite the mixed results shown in country-by-country results, we conclude in favor of the presence of panel unit root and stationarity of the first differences; I(1,1).

As a next step, the results of panel cointegration tests support the view that there is cointegration between FDI stock and quality of institutions (not posted). Pedroni's panel cointegration test is a one-sided test. The test statistics asymptotically follow normal distribution and have a critical value of –1.64 except v-statistics. From our result, ADF group statistics, the most powerful statistics for data with short period, is –2.847, are in agreement with the presence of co-integration. Thus, we interpret the result as the evidence of the existence of a long-run relationship.

4.2. Long Run Elasticity

Appendix 1 reports panel FMOLS result as well as country-by-country estimates. As a unit root test result, the country-by-country estimates show mixed results implying that cointegration is highly varied across countries. For some countries, both FDI and quality of institutions are significant and positively linked, while this is not the case with others. For 11 countries, FDI has a negative and significant effect on quality of institutions. Five of these countries are Latin American; the rest of them are African except Taiwan. What is interesting is that each of these 11 countries experienced a negative feedback effect of quality of institutions on FDI and six countries follow I (1,1). Negative coefficients may reflect the magnified adverse effect of FDI on quality of institutions in highly corrupt environments as suggested by Hellman et al. (2002). However, both FDI and institutional index for these countries are around mean values. So there may be other reasons, such as structural break, behind the dispersion of statistics across countries. For panel level, the coefficient of institutions is 0.15 with a t-statistic of 3.73. FDI stock has a negative (-0.04) but marginally significant effect (1.85) on institutions at 10% level. This result is still consistent with the presence of cointegration, but the reversed sign of institutional effect needs to be questioned considering the fact that still more countries in the sample show the positive cointegrating relationship, although the short time-series cointegration test is not very reliable. Thus, we separately test for subsamples with I (1,1). 26 countries follow I (1,1) around level, and 33 countries have I (1,1) with trend. The result of panel FMOLS for these countries are reported in Appendix 3. There is strong and significant evidence for positive bi-directional causality for subsamples. From this example, nonstationarity of countries seems to play an important role in determining the cointegrating relationship for an average country in the sample. We reinvestigate this issue in the next section using the test results from the error correction model using system GMM estimator.

4.3. Error Correction Model

Table 5 presents the results of two-step system GMM estimation for the panel error correction model. When the current change of FDI is taken as a dependent variable, the one-step estimator posted on the bottom row shows that long-run elasticity is positive and significant between FDI and institutional quality. Also, the coefficient of error correction term is negative and statistically significant. We see a similar result for the equation of differenced institutional quality. The long-run elasticity has a significant and positive sign. The error correction term shows that short-term change of quality of institutions responds to the deviation from long-run equilibrium. These results confirm the evidence of a co-integrating relationship between two variables. As to short-run dynamics, lagged FDI in first difference is negative and significantly related to contemporaneous change of FDI.

However, there is no evidence of short-term causality from institutional quality. The current change of institutional variables affects differenced FDI positively while the lagged one has negative effect, but both effects are statistically insignificant. When the dependent variable is differenced value of quality of institutions, lagged change of its own values has a positive and significant effect on current change. Differenced contemporaneous FDI positively affects change of institutions, while lagged FDI in first difference enters negatively. The coefficients of both variables are insignificant. To sum up, Table 5 supports the presence of long-run bi-directional cointegration, while it shows no evidence in favor of short-term causality between FDI and quality of institutions. The result of specification tests are satisfactory. The *p*-values of Sargan tests agree with the validity of instrumental variables and there is no evidence of second order serial correlation for two equations.

Estimation results of ECM with one-step system GMM estimator are reported in Table 6. The magnitudes of short-term coefficients are not consistent with those in Table 5. Moreover, change of institutional variables in contemporaneous values and current change of FDI have opposite sign to that of the two-step result. However, this may not make a critical difference in that all the short-run dynamics turn out to be insignificant again. Table 6 shows the same result for long-run causality. The significantly negative coefficient of error correction term confirms the existence of long-run co-integration between the quality of institutions and FDI. The second column, representing the impacts on first-differenced institutional variables, shows the same results for the case of the change of FDI. Thus, despite the gap in coefficients of short-run dynamics, these results from one-step system GMM give us the same implication as the two-step estimator with evidence of long-run cointegration and no evidence of short-run causality. Again, the data pass the specification test. We interpret this result as being consistent with the result from two-step system GMM estimator.

5. Conclusions

There seems to be consensus on the argument that quality of institutions of a host country is one of the major factors in attracting foreign direct investment. What is less known is the two-way causality between FDI and quality of institutions. In this paper, we examine causal relationships between foreign direct investment and institutions. An error correction model is estimated using system GMM estimation methods followed by a unit root test and a cointegration test. Our empirical findings suggest that there is a long-run relationship between two variables and the causality is bi-directional. However, there is no clear evidence in favor of short-run causality between institutional quality and FDI.

These results suggest that previous literature should be more cautious in placing too much emphasis on the view that policies aimed at establishing high standards of institutions as a prerequisite would lead to more inward FDI in developing countries, since the effect may not be apparent in the short period. Rather, the policy should aim at long-term improvement of institutional quality. Also the role of FDI inflow should not be overlooked because it can lead to permanent changes in institutions.

For future work, we can explore the role of other determinants such as market potential, productivity or infrastructure of a host country in FDI decisions or institutional quality in multivariate framework, while we focused only on bivariate causality in this paper. Furthermore, additional work could be done to take account of possible structural breaks for both variables. Particularly for FDI, this problem was well documented in UNCTAD (2005). As recording practices differ across countries and change over time, FDI time series data have structural breaks, though it is unlikely that a common regime change occurs in all countries. It could be a potential obstacle to yield more reliable unit root test results.

References

Aizenman, J. and Spiegel, M. (2002), "Institutional Efficiency, Monitoring Costs, and the Investment Share of FDI", *NBER Working Paper No.9324*

Alfaro, L., S. Kalelmi-Ozcan and V. Volosovych (2005), "Capital Flows in a Globalized World: The Role of Policies and Institutions", in *Capital Controls and Capital Flows in Emerging Economies: Policies, Practices and Consequences.* NBER, *forthcoming*.

Arellano, M. and O. Bover (1995), "Another look at the instrumental variable estimation of error-components models", *Journal of Econometrics*, 68, 29-52

Banerjee, A., J. Dolado and R. Mestre (1998), "Error-Correction Mechanism Tests for Cointegration in a Single-Equation Framework", *Journal of Time Series Analysis*, 19, No.3

Banerjee, A., J. Galbraith and J. Dolado (1990), "Dynamic Specification and Linear Transformations of the Autoregressive-Distributed Lag Model", *Oxford Bulletin of Economics and Statistics*, 52, 1

Blundell, R. and Bond, S. (1998), "Initial Conditions and Moment Restrictions in Dynamic Panel Data Models", *Journal of Econometrics*, 87, 115-143

Bond, S., A. Hoeffler and J. Temple (2001), "GMM Estimation of Empirical Growth Models", *Oxford University*, memeo

Campos, N. and Kinoshita, Y. (2003), "Why Does FDI Go Where it Goes? New Evidence from the Transition Economies", *IMF Working Paper* No.228

Carkovic,M. and Levine, R. (2002), "Does Foreign Direct Investment Accelerate Economic Growth?", *University of Minesota Dept. of Finance Working Paper*

Christopoulos, D. and Tsionas, E. (2004), "Financial Development and Economic Growth:

Evidence from Panel Unit Root and Cointegration Tests”, *Journal of Development Economics,* 73, 55-74

Engle, R. and Granger, C. (1987), “Cointegration and Erro-Correction: Representation, Estimation, and Testing”, *Econometrica*, 55, 251-276

Gastanaga, V.M., J. Nugent and Pashamova, B. (1998) “Host country reforms and FDI inflows: How much difference do they make?”, *World Development*, 26 (7), 1299-1314

Hadri, K. and Larsson, R. (2005), “Testing for Stationarity in Heterogeneous Panel Data Where the Time Dimension is Finite”, *The Econometrics Journal*, 8, No. 1, 55-69

Hausmann, R. and Rodrik, D. (2003), “Economic Development as self-discovery”, *Journal of Development Economics*, 72, 603-633

Heins, J.R. Jr. (1995), “Forbidden Payment: Foreign Bribery and American Business After 1977”, *NBER Working Paper* No.5266

Hellman, J., Jones G., Kaufmann D. (2002), “Far From Home: Do Foreign Investors Import Higher Standards of Governance in Transition Economies?”, *World Bank Working Paper*

Im, K., Pesaran, H. and Y. Shin (1995), “Testing for Unit Roots in Heterogeneous Panels”, manuscript, *University of Cambridge*

International Country Risk Guide (2005), *PRG group*

Kang, S. (2004), “Overseas Entry Decision and Ownership Strategy of Japanese Companies: Institution and Corporate Governance”, *Institute of Policy and Planning Sciences.Discussion Paper Series No.1035, University of Tsukuba,* May, 1-37.

Kwiatkowski, D., P. Phillips, P.Schmidt and Y. Shin (1992), “Testing the Null of Hypothesis of Stationarity Against the Alternative of a Unit Root”, *Journal of*

Econometrics, 54, 159-178

Larrain, F., and Tavares, J. (2004), “Does Foreign Direct Investment Decrease Corruption?”, *Cuadernos De Economia*, 41, 217-230

Lee, J. and Mansfield, E. (1996), “Intellectual property protection and U.S. foreign direct investment”, *Review of Economics and Statistics*, 78, 181-186

North, D. (1990), “Institutions, Institutional Change and Economic Performance”, *Cambridge University Press*

Pedroni, P. (1999a), “Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors”, *Oxford Bulletin of Economics and Statistics*, Special Issue (1999)

Pedroni, P. (1999b), “Fully Modified OLS for Heterogeneous Cointegrated Panels”, *Indiana University Working Paper*, memeo

UNCTAD (2005), “Transnational Corporations and the Internationalization of R&D”, *World Investment Report*

Wei, S-J. (2000), “How Taxing is corruption on International Investors?”, *The Review of Economics and Statistics*, 82(1), 1-11

Yasar, M., C. Nelson and R. Rejesus (2004), “ The Dynamics of Exports and Productivity at the Plant Level: A Panel Data Error Correction Model (ECM) Approach”, *Emory University*, memeo

Zhang, X. and Fan, S. (2001), “How Productive is Infrastructure? New Approach and Evidence from Rural India”, *EPTD Discussion paper* No.84

Table 1

Summary Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
FDI	1240	2.4409	1.335	-2.279	7.492
Institutions	1240	-0.1546	0.822	-3.135	2.39

Notes: FDI is log of FDI stock per capita. Institutional quality is transformed using inverse logit function of normalized scores indicating quality of each subcomponents-corruption, law and order, and bureaucracy quality.

Table 2

IPS Panel unit root tests

Variables	Trend		No Trend	
	Levels	First differences	Levels	First differences
FDI	**−2.12**	-3.906***	**−1.043**	-3.774***
Institutions	**−1.536**	-3.537***	**−1.390**	-3.477***

Notes: Boldface values denote sampling evidence in favor of unit roots. *** represents rejection of the null of unit roots at the l% level of significance.

Table 3

KPSS country-by-country stationarity tests (number of countries)

Variables	Trend	No Trend	
	Levels	Levels	First differences
FDI	39	50	8
Institutions	41	34	10

Notes: The entries of table denote number of countries with sampling evidence in favor of unit roots.

Table 4

KPSS Panel Stationarity tests (Lag=2): Hadri and Larsson

Variables	Trend	No Trend	
	Levels	Levels	First differences
FDI	**11.155*****	**21.83*****	1.87
Institutions	**10.927*****	**14.217*****	1.85

Notes: Boldface values denote sampling evidence in favor of unit roots. *** represents rejection of the null of stationarity at the l% level of significance.

Table 5

Panel Error Correction Model (Two Step System GMM)

Explanatory variables	Dependent variables	
	ΔFDI_t	$\Delta Inst_t$
$\Delta Inst_t$	0.443 (0.52)	
$\Delta Inst_{t-1}$	-0.14 (0.819)	0.346 (0.000)
ΔFDI_t		-0.014 (0.773)
ΔFDI_{t-1}	-0.045 (0.392)	-0.062 (0.208)
$FDI_{t-2} - \phi Inst_{t-2}$	-0.05 (0.000)	
$Inst_{t-2} - \phi FDI_{t-2}$		-0.069 (0.000)
Constant	0.029 (0.379)	0.008 (0.277)
Number of Observation	1116	1116
Sargan Test	0.435	0.219
AR (1) in first differences	0.000	0.000
AR (2) in first differences	0.577	0.40
Long run Elasticities	0.31 (0.000)	0.295 (0.000)

Notes: Time-dummies are included for equation of change of FDI. *P*-values are in parentheses.

Table 6

Panel Error Correction Model (One Step System GMM estimator)

Explanatory variables	Dependent variables	
	ΔFDI	$\Delta Inst$
$\Delta Inst_t$	-0.032 (0.404)	
$\Delta Inst_{t-1}$	-0.009 (0.798)	0.315 (0.000)
$Inst_{t-2} - FDI_{t-2}$		-0.096 (0.000)
$Inst_{t-2}$	-0.051 (0.014)	
ΔFDI_t		-0.021 (0.475)
ΔFDI_{t-1}	0.118 (0.074)	-0.042 (0.153)
$FDI_{t-2} - Inst_{t-2}$	-0.0216 (0.012)	
FDI_{t-2}		-0.109 (0.000)
constant	0.123 (0.002)	0.016 (0.748)
Number of Observation	1116	1116
Sargan test(P-value)	0.212	0.465
AR(1) (P-value)	0.000	0.000
AR(2) (P-value)	0.313	0.458

Notes: Time-dummies are included for equation of change of FDI. *P*-values are in parentheses.

Appendix 1
Fully Modified OLS

	Individual FMOLS results			
	FDI		Institutional Quality	
	Coefficient	t-statistic	Coefficient	t-statistic
Algeria	-0.60	(-1.12)	-0.40	(-3.09)
Angola	-2.13	(-0.62)	-0.13	(-2.16)
Argentina	1.85	(1.86)	0.02	(0.16)
Bahrain	-0.08	(-0.09)	0.01	(0.05)
Bangladesh	0.25	(1.25)	0.39	(0.65)
Bolivia	0.60	(4.13)	0.87	(3.49)
Brazil	-0.74	(-2.78)	-0.31	(-0.91)
Chile	1.20	(4.87)	0.60	(4.26)
China	0.15	(0.13)	0.12	(0.95)
Colombia	-1.36	(-6.21)	-0.54	(-6.98)
Costa Rica	-0.50	(-2.05)	-0.82	(-4.66)
Cote d'Ivoire	-0.72	(-5.41)	-1.11	(-6.69)
Dominican Rep.	-0.23	(-0.35)	-0.30	(-2.00)
Ecuador	-2.56	(-3.08)	-0.21	(-3.56)
El Salvador	0.58	(2.35)	0.49	(1.43)
Ghana	-0.38	(-0.82)	-0.04	(-0.12)
Guatemala	-0.05	(-0.60)	-1.20	(-0.78)
Guinea	5.27	(9.23)	0.14	(8.80)
Haiti	-0.13	(-1.43)	-0.63	(-0.54)
Honduras	1.73	(10.55)	0.46	(11.85)
Hong Kong	-0.18	(-0.63)	-0.33	(-0.77)
India	1.72	(1.86)	0.16	(1.99)
Indonesia	-0.09	(-1.06)	-1.19	(-0.82)
Iran	0.11	(0.73)	-0.02	(-0.05)
Jamaica	0.24	(0.55)	0.18	(0.59)
Jordan	-0.21	(-0.57)	-0.19	-0.44)
Kenya	-0.10	(-0.87)	-0.04	(-0.03)
Kuwait	0.98	(1.35)	0.20	(1.79)
Lebanon	0.30	(0.64)	0.04	(0.15)
Liberia	0.07	(0.13)	0.08	(0.29)

Malawi	0.68	(2.17)	0.33	(1.33)
Malaysia	-0.08	(-0.18)	-0.20	(-0.62)
Mali	1.87	(2.29)	0.12	(1.24)
Mexico	-0.50	(-0.55)	-0.07	(-0.58)
Nicaragua	-4.54	(-2.07)	0.04	(0.80)
Nigeria	0.19	(0.56)	0.62	(1.67)
Pakistan	1.46	(3.32)	0.41	(5.18)
Panama	0.26	(2.56)	1.32	(1.53)
Paraguay	0.14	(1.10)	0.67	(0.62)
Peru	1.48	(6.17)	0.48	(5.01)
Philippines	0.20	(2.13)	1.54	(1.82)
Saudi Arabia	1.24	(2.32)	0.30	(2.47)
Senegal	-1.10	(-0.59)	-0.11	(-2.01)
Singapore	0.43	(2.02)	0.42	(1.03)
South Africa	-0.67	(-8.78)	-1.31	(-9.98)
Korea	-0.16	(-0.49)	-0.41	(-0.83)
Sri Lanka	0.43	(2.86)	1.05	(2.48)
Sudan	1.03	(0.54)	0.03	(0.48)
Syria	0.84	(2.15)	0.29	(1.80)
Taiwan	-0.65	(-4.49)	-0.79	(-3.38)
Tanzania	0.56	(0.47)	-0.30	(-1.42)
Thailand	-0.24	(-0.46)	-0.19	(-1.49)
Togo	-0.95	(-1.65)	0.28	(3.25)
Trinidad and Tobago	1.60	(3.32)	-3.07	(-1.52)
Tunisia	-0.02	(-0.53)	-0.06	(-0.08)
Turkey	0.06	(0.43)	0.06	(0.32)
UAE	-0.10	(-0.18)	0.28	(4.96)
Uganda	2.10	(4.58)	-0.28	(-1.77)
Uruguay	-0.74	(-1.00)	-0.35	(-3.21)
Venezuela	-0.98	(-1.88)	0.60	(3.20)
Zambia	0.61	(2.27)	-0.35	(-1.29)
Zimbabwe	0.02	(0.05)	-0.35	(-1.29)
Panel FMOLS	**0.15**	**(3.73)**	**-0.04**	**(1.85)**

Notes: t-values are in parentheses.

Appendix 2

Quantiles of KPSS statistics

Nominal values		**FDI**		**Quality of Institutions**	
		(Lag=1)	(Lag=2)	(Lag=1)	(Lag=2)
90%	0.349	0.411	0.390	0.434	0.407
95%	0.446	0.461	0.420	0.452	0.499
99%	0.762	0.601	0.526	0.606	0.561

Notes: FDI and Quality of institutions are values in first differences.

Appendix 3

Panel FMOLS for countries with I(1,1)

	FDI		Quality of Institutions	
	Coefficient	t-statistics	Coefficient	t-statistics
I(1,1) around level	0.88	**10.47**	0.38	**8.51**
I(1,1) around trend	0.29	**3.61**	0.17	**2.78**

Appendix 4

Countries with negative signs for both direction of causality

I(1,1) around level	I(1,1) trend
Colombia	Algeria
Cote d'Ivore	Costa Rica
Ecuador	South Africa
Venezuela	Taiwan
	Togo

Figure 1

Histogram of individual KPSS stat.

Differenced FDI (Lag=1)

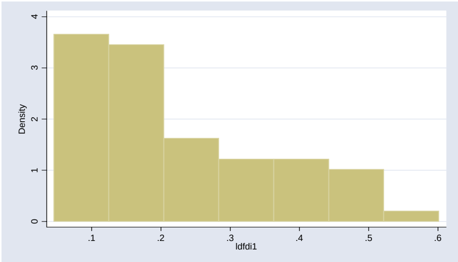

Figure 2

Histogram of individual KPSS stat.

Differenced FDI (Lag=2)

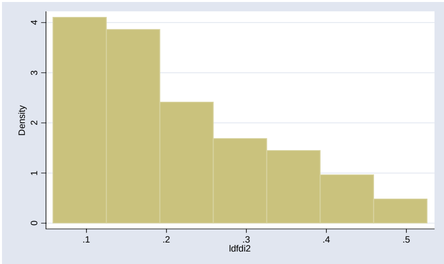

Figure 3

Histogram of individual KPSS stat.

Differenced Quality of institutions (Lag=1)

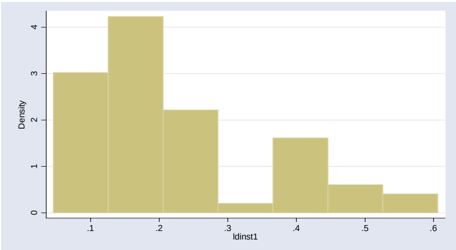

Figure 4

Histogram of individual KPSS stat.

Differenced Quality of institutions (Lag=2)