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Analysis of FDI Inflows into China from ASEAN-5 Countries: A Panel Cointegration Approach

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This study analyses the foreign direct investment (*FDI*) inflows into China from ASEAN–5 countries using the panel cointegration approach. The *FDI* model has been utilized in determining factors that influence *FDI* inflows into China from ASEAN–5 countries, namely, Malaysia, Thailand, the Philippines, Indonesia and Singapore. Variables such as trade openness (*OPENNESS*), exchange rate of China relative to each of individual ASEAN–5 countries (*RELEXC*), fixed capital formation (*FCF*), and gross domestic product (*GDP*) are used for the period of 1990 – 2004. The empirical results indicate that for most countries *OPENNESS* and *GDP* are significant variables in explaining the flow of *FDI* to China. Meanwhile, *FCF* is only significant for Malaysia. Conversely, *RELEXC* is not statistically significant for all countries. It is hoped that this finding can be used by researchers and policy makers in making decision on multilateral relationship between China and ASEAN–5 countries.

1. Introduction

Foreign direct investment (*FDI*) plays an important role in the growth and development of not only the developed countries but also in the developing countries. Besides the capital that it brought in, it also introduces new and modern technology which provides market opportunities and linkages to export. Countries are competing with each other to offer a lucrative incentive plans to attract *FDI*.

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The Asian region has always been considered as a prudent centre for investment especially from the United States of America (USA), Japan, United Kingdom (UK) and other European countries. Globalization and integration of economic activities across the world forced the government of the Asian countries to attract *FDI* which later on translated into rapid growth in these economies. Asian countries are implementing new plans and policies to attract more and more *FDI* which will bring in new innovation and automation based technologies that can rejuvenate the host country's existing manufacturing base. Furthermore, the Asian region attracts *FDI* inflows as a result of her abundant natural resources, highly skilled, experienced and knowledge-versed labour, and huge size of domestic market.

China has been opening up its economy for more than twenty years; however its accession to World Trade Organization (WTO) on 11 December 2001 implies extensive consequences for its economy. China's opening up policy has aimed at promoting exports, while protecting the domestic market. This was achieved through a dualistic trade regime, which has granted tariff exemptions on imports of intermediate goods by export–oriented industries, and through a selective policy, which has channelled *FDI* into manufacturing production targeted for exports or for import substitution. As a result, *FDI* inflows have played a major part in the opening up of China's industry and its integration into the international division of labour. The rapid expansion of its international trade and large capital inflows provide evidence of the increasing integration of China in the world economy. Since 1980, China's share in international trade has trebled, rising from less than one percent to more than three percent in 1999. During the first 6 months of 2012, China has surpassed the USA as the world's largest recipient of global FDI with a total of USD59 billion compared to *FDI* flowing to the USA totalling USD57.4 billion

The cooperation and partnership between China and ASEAN has long been established through various channels for attaining certain goals. One of the major channels is *FDI* inflows. According to Shu and Zeng (2006), *FDI* inflows from ASEAN–5 into China in 2004 was about fifty times as much as it was in 1990 (see Table 1). During 1994 – 2004, the cumulative amount of China's actually utilized *FDI* from ASEAN–5 reached USD33.73 billion, which exceeded the cumulative amount of China's actually utilized *FDI* from the UK, France and Germany combined, which was USD27.21 billion. Based on country, Singapore recorded the highest amount of *FDI* inflows to China and then followed by Malaysia. This fact is in line with the argument by Ellingsen *et al.* (2006) that Singapore is one of the most important outward investors in the developed countries. Singaporean direct investors were strongly encouraged by the government to reach beyond ASEAN and were increasingly shifting their attention to other Asian host countries, particularly China. Furthermore, the growth of *FDI* inflows from ASEAN–5 into China increased tremendously in 1992 and 1993, but showed a declining trend after the year 1994. The worst *FDI* growth recorded in 1999 that due to East Asian financial crisis in the middle of 1997. This largely unforeseen crisis and its aftermath caused deterioration in the macroeconomic fundamentals, particularly slower economic growth of Thailand, Malaysia and Indonesia. As a result, these countries reduced their *FDI* inflows into China. This gives rise to the issues of whether the ASEAN–5 countries are able to maintain investment in China and the emerging factors that significantly affect the ability of ASEAN–5 countries invest in China.

Table 1: Foreign Direct Investment in China by ASEAN–5 Countries, 1990 – 2004 (Amount contracted in USD million)

Year	Country					ASEAN–5	Growth (%)
	Malaysia	Indonesia	Singapore	Thailand	The Philippines		
1990	64	100	5043	672	167	6046	
1991	196	218	5821	1962	585	8782	45.25
1992	2467	2017	12593	8432	1655	28390	223.27
1993	9142	6575	49180	23437	12250	102385	260.64
1994	20099	11570	11791	23487	14040	189300	84.89
1995	25900	11163	186061	28824	10578	265356	40.18
1996	45995	9354	224716	32818	5551	319396	20.37
1997	38183	7998	260641	19400	15563	342800	7.33
1998	34049	6879	340397	20538	17927	422318	23.20
1999	23771	12917	264249	14832	11728	328877	-22.13
2000	20288	14694	217220	20357	11112	284458	-13.51
2001	26298	15964	214355	19421	20939	298395	4.90
2002	36786	12164	233720	18772	18600	325594	9.12
2003	25103	15013	205840	17352	22001	292543	-10.15
2004	38504	10452	200814	17828	23324	304053	3.93

Sources: Shu and Zeng (2006).

Given the above scenario, we are encouraged to undertake further analysis to identify factors that significantly affect the *FDI* inflows into China from the selected ASEAN–5 countries. In this article, we postulate that China's macroeconomic fundamentals can influence *FDI* inflows from the ASEAN–5 countries.

Furthermore, this study is also motivated by the fact that there has been little econometric modeling and evidences of how the macro fundamentals of China affect the inflows of capital from the ASEAN–5 countries. This paper tries to fill this gap by analyzing the link between China and the ASEAN–5 countries through *FDI* inflows activities. This link has never been addressed aggressively in the literature particularly after WTO accession protocol. Therefore, this study aims to provide new evidence of macro-determinants of *FDI* inflows between China and ASEAN–5. Since China has become very important in the world economic operation and an attractive site for *FDI* (Naughton, 1996) after admission in WTO, ASEAN–5 countries would gain some significant benefits through *FDI* activities in China.

The paper is organized as follows. Section 2 reviews the existing literature on determinants of *FDI* inflows. Section 3 describes the data used and the methodology of determining factors that influence the level of *FDI* inflows into China. The empirical results of the study are reported in Section 4 and Section 5 presents the conclusion.

2. Literature Review

Many studies have been carried out to analyze the determinants of *FDI* and its effect to host countries, particularly developing countries. As stated in many studies, *FDI* plays an important role in these economies which are generally lacking in terms of technology as well as capital to fund the projects (Borenzstein *et al*., 1998; Estrin *et al*., 2000; Buckley *et al*., 2001). In addition, Calvo *et al*. (1996) and Moreno (2000) stated that capital flows from one country to another also allow investors to diversify their risks.

As the time progress and because of many factors, the trend of *FDI* inflows is not just from developed countries to developing or under developing countries, but the reverse or outward *FDI* may exist. This phenomenon also happens among ASEAN–5 countries. Even though, these countries are considered as developing countries, they intend to invest in abroad in order to grab the opportunities and advantageous offered by the recipient countries.

One of the specific studies related to *FDI* inflows in China and ASEAN has done by Shu and Zeng (2006). Their descriptive analyses have indicated three factors that exert significant influence on *FDI* inflows between China and ASEAN. These factors are new bilateral economic agreements, China's new mega economic zone, and ASEAN reforms and new foreign policy. Bilateral agreements such as Framework Agreement on Comprehensive Economic Cooperation (FACEC) between ASEAN and People's Republic of China signed in 2002, for instance, can strengthen economic relations, which is a strategic goal of both sides. A large variety of Chinese products have been exported, on a large scale, into ASEAN member countries. Following the FACEC agreement, China provides survival and growth opportunities for ASEAN investors and traders.

There are many emerging microeconomic and macroeconomic fundamentals that influence a country to invest abroad or to receive *FDI* inflows from abroad. However, this brief review gives more emphasize to the effects of macroeconomic fundamentals on *FDI* inflows. In addition to geographic location, the important macroeconomic fundamentals of host/recipient countries such as gross domestic product (*GDP*), exchange rates, trade openness, and capital accumulation are considered as important determinants.

Openness of one location is one of the traditional variables to explain *FDI* movements. It is defined as the ratio of total trade (export plus import) to *GDP*. MNEs engaged in export oriented investment prefer to locate in a more open economy as increased imperfections that accompany trade protection generally imply higher transaction costs associated with exporting. A study done by the Organization for Economic Cooperation and Development (2000) find that trade openness is one of the main determinants of *FDI* in China. China has adopted the 'export promotion strategy' where it has implemented economic reforms and open-door policies, at the same time it also promotes trade by concluding several bilateral trade arrangements. This can be seen in 1990s where there has been a substantial progress in reducing tariff barriers, i.e. the average tariff rate on imports declined from 42.9 percent in 1992 to 17.6 percent in 1997. With regard to this variable, the expected impact of the degree of openness on *FDI* can be mixed; it can attract the foreign capital to the host area and also can increase competition between the foreign and domestic firms on it. Meanwhile, Acemoglu and Zilibotti (1997) explain that greater openness in international transactions helps direct financial flows from capital-abundant towards capital-scarce countries. Adhikary (2011) has claimed that the degree of trade openness is likely to influence the flows of international capital in terms of risk-return relationship.

Exchange rate has greater impact on the flow of foreign capital into host country [see Froot and Jeremy (1991), Pain and Van Welsum (2003), Philips *et al.* (2008), Gottschalk *et al.* (2008), Baek and Okawa (2001), Nyarko and Nketiah-Amponsah (2011), Chaudary *et al.* (2012), Ullah *et al.* (2012), and Ngowani (2012)]. Chaudary *et al.* (2012), for instance, have studied the effects of exchange rate on *FDI* in the Asian economies; with Pakistan, India, Sri Lanka, and Bangladesh representing South Asian. Malaysia, Indonesia, Singapore and Thailand representing the Southeast Asian region; China, Japan and South Korea representing the country in East Asia. While Turkey, Iran, and Israel representing West Asia. They have used Autoregressive Distributed Lag (ARDL), cointegration and error correction model (ECM) for performing empirical analysis. They found that more than half of the selected countries show no significant relationship between *FDI* and exchange rate. Effect of short-term and long-term exists in Pakistan, India, Sri Lanka, South Korea, Turkey and Israel. The results indicated there was no relationship between *FDI* and exchange rates for Bangladesh, China, Malaysia, Indonesia, Thailand, Singapore, and Iran.

Furthermore, economic growth of the host country is another main determinant of *FDI* inflows. This variable is used in many studies such as Toda and Yamamoto (1995), Wang and Swain (1995), Zhang (1999, 2000, 2002), Barthelemy and Demurger (2000), Wei and Liu (2001), Jordan (2002), Ali and Guo (2005), and Le and Liu (2005). They agreed that one of the main factors is the growth of the host country which can be measured by the *GDP*, *GDP* per capita, gross national product (*GNP*), and *GNP* per capita. Thus, the variable *GDP* is expected to have a positive significant relationship with *FDI* inflows. Borensztein *et al.* (1998), for instance, have argued that *FDI* inflows are positively related to per capita *GDP* growth provided the host country has a highly educated workforce. Meanwhile, Chowdhury and Mavrotas (2003) find unidirectional causality running from growth to *FDI* in the case of Chile but find bidirectional causality for Thailand and Malaysia.

The earlier growth models by Harrod (1939) and Domar (1946) explained that capital formation raises the standard of living, which in turn results in higher growth. Then, Solow (1956) argues that capital formation increases labor productivity in a dynamic process of investment growth and finally enhancing economic growth. These evidences recognize the role of capital formation in economic growth and then create a basis for attracting *FDI* inflows into a particular country. Analytically, capital formation can improve domestic investment through positive spillovers and by creating complementary industries. Thus, capital formation has a dynamic effect on *FDI* inflows. This is supported by Adhikary (2011) where he mentioned that the level of capital formation is likely to influence *FDI* and economic growth. Neo-classical growth model postulates that developing economies that have a lower initial level of capital stock tend to have higher marginal rate of returns (productivity) and growth rates if adequate capital stock is injected. In other words, in a capital shortage economy, the marginal productivity of investment is increased in the short-run when additional capital is injected in the form of long-term investment like *FDI*, and this increased productivity influences economic growth in the long-run.

At the macro level, by and large, previous literature suggests that macroeconomic determinants contribute to *FDI* inflows through either direct or indirect FDI. Direct FDI is a type of investment where investment activities are carried out by the foreign-owned firm such as parent (headquarter) directly. Meanwhile, indirect FDI is a type of FDI by foreign affiliate. Investment activities are carried out by firm, which themselves are affiliates of foreign Multinational Enterprises (MNEs) such as a regional headquarter.

3. Methodology

3.1 Data

Secondary data are used in the study. Due to short time spans of time series data, we have pooled cross–section and time series data to form a balanced panel data. Using the short time span of time series data may yield unreliable results. The balanced panel consists of annual data for *FDI* inflows into China from five selected ASEAN–5 countries, namely Malaysia, Thailand, The Philippines, Indonesia and Singapore for the period of 1990 – 2004 and data of each variable is measured in US dollars. The use of panel data allows us to study both the changes in the independent variables of a single ASEAN–5 country over time and the variation in these variables of many ASEAN–5 countries at a given point in time. The data are gathered and verified from various sources i.e. International Financial Statistics by International Monetary Fund (IMF), Direction of Trade Statistics, World Development Indicators and World Debt Tables.

3.2 Model Specification

Based on previous literature, *FDI* inflows are mainly determined by trade openness (*OPENESS*), exchange rate (*RELEXC*), fixed capital formation (*FCF*), and *GDP*. It is described by the basic function as Equation [1]. Furthermore, the empirical model form for this function is given by Equation [2].

$$FDI = f(OPENESS, RELEXC, FCF, GDP) \qquad [1]$$

$$FDI_{it} = \beta_0 + \beta_1 OPENESS_{it} + \beta_2 RELEXC_{it} + \beta_3 FCF_{it} + \beta_4 GDP_{it} + \varepsilon_{it} \qquad [2]$$

In Equation [2], *FDI* represents the total *FDI* inflows into China from ASEAN–5 countries, *OPENESS* represents the level of trade openness and it is measured by the trade-intensity ratio, which is the share of export and import in *GDP*, *RELEXC* is the exchange rate of China relative to each of the individual ASEAN–5 countries, *FCF* is the total of fixed capital formation in China, and *GDP* is the gross domestic product of China. The choice of these variables relies on the data accessibility and the value of data in dollar America. The β_0 is a constant term and β_1 to β_4 are estimated parameters in the model and i is a cross–section data for countries referred to, and t is a time series data and ε_{it} is an error term. Both the coefficients, β_3 and β_4, are expected to carry a positive sign.

3.3 Estimation Procedure

Estimation procedure of the relationship between *FDI* inflows into China and independent variables in Equation [2] involves two main steps, cointegration analysis and coefficient estimation.

3.3.1 Cointegration Analysis

We have employed panel cointegration approach to perform cointegration analysis. This method is used to investigate the existence of the long run cointegration among the variables. In order to investigate the possibility of panel cointegration, it is first necessary to determine the existence of unit roots in the data series. For this study, we have chosen the Im, Pesaran and Shin (IPS, hereafter) (1997), which is based on the well-known Dickey–Fuller procedure. IPS proposed a test for the presence of unit roots in panels that combines information from the time series dimension with that from the cross section dimension, such that fewer time observations are required for the test to have power. The use of this method is motivated by that this is commonly used and superior test power (Chou and Lee, 2003). Many economic researchers such as Lee *et al.* (1997), Sarantis and Steward (1999), Canzoneri *et al.* (1999), and Chou and Lee (2003) have also applied this method in their analysis of the long–run relationships in panel data.

IPS begins by specifying a separate ADF regression for each cross-section with individual effects and no time trend:

$$\Delta y_{it} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \qquad [3]$$

where $i = 1, \ldots, N$ and $t = 1, \ldots, T$

IPS use separate unit root tests for the N cross–section units. Their test is based on the Augmented Dickey – Fuller (ADF) statistics averaged across groups. After estimating the separate ADF regressions, the average of the t – statistics for p_1 from the individual ADF regressions, $t_{iT_i}(p_i)$:

$$\bar{t}_{NT} = \frac{1}{N}\sum_{i=1}^{N} t_{iT}\left(p_i \beta_i\right) \qquad [4]$$

The t – bar $(\bar{t})$ is then standardized and it is shown that the standardized $\bar{t}$ statistic converges to the standard normal distribution as N and $T \rightarrow \infty$. IPS (1997) showed that $\bar{t}$ test has better performance when N and T are small. They proposed a cross–sectionally demeaned version of both test to be used in the case where the errors in different regressions contain a common time-specific component.

The next step is to test for the existence of a long–run cointegration between *FDI* and the independent variables using panel cointegration tests suggested by Pedroni (1999, 2004). Pedroni (1999) determined that the tests are appropriate to estimate residuals from a cointegration regression after normalizing the panel statistics with correction terms. The test procedures proposed by Pedroni make use of estimated residual from the hypothesized long–run regression of the following form:

$$y_{i,t} = \alpha_i + \delta_i t + \beta_{1i,t} x_{1i,t} + \beta_{2i} x_{2i,t} + \ldots + \beta_{Mi} x_{Mi,t} + e_{i,t} \qquad [5]$$

for $t = 1,\ldots,T;\ i = 1,\ldots,N;\ m = 1,\ \ldots,\ M,$

where T is the number of observations over time, N number of cross-sectional units in the panel, and M number of regressors. In this set up, α_i is the member specific intercept or fixed effects parameter which varies across individual cross-sectional units. The same is true of the slope coefficients and member specific time effects, $\delta_i t$.

Pedroni (1999, 2004) proposes the heterogeneous panel and heterogeneous group mean panel test statistics to test for panel cointegration. He defines two sets of statistics. The first set of three statistics $Z_{\hat{v},N,T}$, $Z_{\hat{\rho}N,T\ 1}$ and $Z_{tN,T}$ is based on pooling the residuals along the within dimension of the panel. The statistics are as follows

$$Z_{\hat{v},N,T} = T^2 N^{3/2}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\,2}\hat{e}_{i,t\ \ 1}^{2}\right) \qquad [6]$$

$$Z_{\hat{\rho}N,T\;\;1} = T\sqrt{N}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\,2}\hat{e}_{i,t\;\;1}^{2}\right)^{1}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\,2}\left(\hat{e}_{i,t\;\;1}\Delta\hat{e}_{i,t}\hat{\lambda}_{i}\right) \qquad [7]$$

$$Z_{tN,T} = \left(\tilde{\sigma}_{N,T}^{2}\sum_{I=1}^{N}\sum_{T=1}^{T}\hat{L}_{11i}^{\,2}\hat{e}_{i,t\;\;1}^{2}\right)^{1/2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\,2}\hat{e}_{i,t\;\;1}^{2}\left(\hat{e}_{i,t\;\;1}\Delta\hat{e}_{i,t}\hat{\lambda}_{i}\right) \quad [8]$$

where $\hat{e}_{i,t\;\;1}$ is the residual vector of the OLS estimation of Equation [5] and where the other terms are properly defined in Pedroni.

The second set of statistics is based on pooling the residuals along the between dimension of the panel. It allows for a heterogeneous autocorrelation parameter across members. The statistics are as follows:

$$\tilde{Z}_{\hat{\rho}N,T\;\;1} = \sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{i,t\;\;1}^{2}\right)^{1}\sum_{t=1}^{T}\left(\hat{e}_{i,t\;\;1}\Delta\hat{e}_{i,t}\hat{\lambda}_{i}\right) \qquad [9]$$

$$\tilde{Z}_{tN,T\;\;1} = \sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{i,t\;\;1}^{2}\right)^{1/2}\sum_{t=1}^{T}\left(\hat{e}_{i,t\;\;1}\Delta\hat{e}_{i,t}\hat{\lambda}_{i}\right) \qquad [10]$$

These statistics compute the group mean of the individual conventional time series statistics. The asymptotic distribution of each of those five statistics can be expressed in the following form:

$$\frac{X_{N,T}\quad\mu\sqrt{N}}{\sqrt{v}} \Rightarrow N(0,1) \qquad [11]$$

where $X_{N,T}$ is the corresponding from of the test statistics, while μ and v are the mean and variance of each test respectively. They are given in Table 2 in Pedroni (1999). Under the alternative hypothesis, Panel v statistics diverges to positive infinity. Therefore, it is a one sided test were large positive values reject the null of no cointegration. The remaining statistics diverge to negative infinity, which means that large negative values reject the null.

3.3.2 Coefficients Estimation

For coefficient estimation, we have adopted the Fully Modified Ordinary Least Squares (FMOLS) procedure based on Christopoulos and Tsionas (2004). In order to obtain asymptotically efficient consistent estimates in panel series, non–exogeneity and serial correlation problems are tackled by employing FMOLS introduced by Pedroni (1996). Since the explanatory variables are cointegrated with a time trend, and thus a long-run equilibrium relationship exists among these variables through the panel unit root test and panel cointegration test, we proceed to estimate the Equation [2] by the method FMOLS for heterogenous cointegrated panels. This methodology allows consistent and efficient estimation of cointegration vector and also addresses the problem of non–stationary regressors, as well as the problem of simultaneity biases. It is well known that Ordinary Least Squares (OLS) estimation yields biased results because the regressors are endogenously determined in the *I*(1) case. The starting point OLS as in the following cointegrated system for panel data:

$$y_{it} = \alpha_i + x'_{it}\beta + e_{it} \qquad [12]$$
$$x_{it} = x_{i,t\ 1} + \varepsilon_{it}$$

where $\xi_{it} = [e_{it}, \varepsilon'_{it}]$ is the stationary with covariance matrix Ω_i. The estimator β will be consistent when the error process $\omega_{it} + [e_{it}, \varepsilon'_{it}]'$ satisfies the assumption of cointegration between y_{it} and x_{it}. The limiting distribution of OLS estimator depends upon nuisance parameters. Following Phillips and Hansen (1990), a semi–parametric correction can be made to the OLS estimator that eliminates the second order bias caused by the fact that the regressors are endogenous. Pedroni (1996, 2000) follows the same principle in the panel data context, and allows for the heterogeneity in the short run dynamics and the fixed effects. FMOLS Pedroni's estimator is constructed as follow:

$$\hat{\beta}_{FM} = \left(\sum_{i=1}^{N} \hat{\Omega}_{22i}^{2} \sum_{t=1}^{T} (x_{it}\hat{x}_t)^2 \right)^{1} \sum_{i=1}^{N} \hat{\Omega}_{11i}^{1} \hat{\Omega}_{22i}^{1} \left(\sum_{t=1}^{T} (x_{it}\bar{x}_t) e_{it} T\hat{\gamma}_i \right) \qquad [13]$$
$$\hat{e}_{it} = e_{it}\hat{\Omega}_{22i}^{1}\hat{\Omega}_{21i}, \qquad \hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^{0}\hat{\Omega}_{22i}^{1}\hat{\Omega}_{21i}\left(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^{0}\right)$$

where the covariance matrix can be decomposed as $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i$ where Ω_i^0 is the contemporaneous covariance matrix, and Γ_i is a weighted sum of autocovariances. Also, $\hat{\Omega}_i^0$ denotes an appropriate estimator of Ω_i^0.

In this study, we employed panel group FMOLS test from Pedroni (1996, 2000). An important advantage of the panel group estimators is that the form in which the data is pooled allows for greater flexibility in the presence of heterogeneity of the cointegrating vectors. Test statistics constructed from the panel group estimators are designed to test the null hypothesis $H_0 : \beta_i = \beta_0$ for all i against the alternative hypothesis $H_A : \beta_i \neq \beta_0$, so that the values for β_i are not constrained to be the same under the alternative hypothesis. Clearly, this is an important advantage for applications such as the present one, because there is no reason to believe that, if the cointegrating slopes are not equal to one, which they necessarily take on some other arbitrary common value. Another advantage of the panel group estimators is that the point estimates have a more useful interpretation in the event that the true cointegrating vectors are heterogeneous. Specifically, point estimates for the panel group estimator can be interpreted as the mean value for the cointegrating vectors (Pedroni, 2001).

4. Empirical Results

Table 2 presents the results of the IPS panel unit root test at level indicating that all variables are *I*(1) in the constant of the panel unit root regression.

These results clearly show that the null hypothesis of a panel unit root in the level of the series cannot be rejected at various lag lengths. We assume that there is no time trend. Therefore, we test for stationarity allowing for a constant plus time trend. In the absence of a constant plus time trend, again we found that the null hypothesis of having panel unit root is generally rejected in all series at level form and various lag lengths. We can conclude that most of the variables are non-stationary in with and without time trend specifications at level by applying the IPS test which is also applied for heterogeneous panel to test the series for the presence of a unit root.

Table 2: Panel Unit Root Test – Im, Pesaran and Shin (IPS)

Variable	Level		First Order Difference	
	Constant	Constant + Trend	Constant	Constant + Trend
FDI	-1.261 (0.104)	-1.043 (0.148)	-2.707** (0.003)	-1.703** (0.044)
OPENESS	-0.918 (0.913)	-1.682 (0.871)	-3.411* (0.000)	-3.506** (0.001)
RELEXC	-1.418 (0.586)	-1.772 (0.821)	-3.183* (0.000)	-3.313** (0.005)
FCF	-2.044 (0.899)	-1.620 (0.899)	-2.451** (0.017)	-1.549*** (0.061)
GDP	-2.010 (0.127)	-1.963 (0.683)	-2.977* (0.001)	-2.887** (0.051)

Note: * and ** indicate rejection of the null hypothesis of no cointegration at one percent and five percent levels of significance

The results of the panel unit root tests confirm that the variables are non–stationary at level. Table 2 also presents the results of the tests at first difference for IPS test in constant and constant plus time trend. We can see that for all series the null hypothesis of unit root test is rejected at both one percent and five percent levels of significance. Hence, based on IPS test, there strong evidence that all the series are in fact integrated of orders one.

We can conclude that the results of panel unit root test (IPS test) reported in Table 2 support the hypothesis of a unit root in all variables across countries, as well as the hypothesis of zero order integration in first differences. At most of one percent significance level, we found that all tests statistics in both with and without trends significantly confirm that all series strongly reject the unit root null. Given the result of IPS test, it is possible to apply panel cointegration method in order to test for the existence of the stable long–run relation among the variables.

The next step is to test whether the variables are cointegrated using Pedroni's (1999, 2001, 2004). This is to investigate whether long–run steady state or cointegration exist among the variables and to confirm what Coiteux and Olivier (2000) state that the panel cointegration tests have much higher testing power than conventional cointegration test. Since the variables are found to be integrated in the same order *I* (1), we continue with the panel cointegration tests proposed by Pedroni (1999, 2001, 2004). Cointegrations are carried out for constant and constant plus time trend and the summary of the results of cointegrations analyses are presented in Table 3.

In constant level, we found that three out of seven statistics reject null by hypothesis of no cointegration at the five percent level of significance for the *adf* – statistic and group ρ – statistic, while the group – *adf* is significant at one percent level. The results of the panel cointegration tests in the model with constant level show that independent variables do hold cointegration in the long run for a group of ASEAN–5 countries with respect to *FDI*. In the panel cointegration test for our model with constant plus trend level, the results indicate that four out of seven statistics reject the null hypothesis of non cointegration at the one percent and five percent level of significance. It is shown that independent variables do hold cointegration in the long run for a group of ASEAN–5 countries with respect to *FDI*. However, since all the statistics conclude in favour of cointegration, and this, combined with the fact that the according to Pedroni (1999) the panel non–parametric (*t* – statistic) and parametric (*adf* – statistic) statistics are more reliable in constant plus time trend, we conclude that there is a long–run cointegration among our variables in ASEAN–5 countries.

Table 3: The Pedroni Panel Cointegration Test

Test	**Constant trend**	**Constant + Trend**
Panel v – Statistic	-0.061	-0.842
Panel ρ – Statistic	-0.988	-1.836**
Panel t – Statistic: (non–parametric)	-1.024	-1.210
Panel t – Statistic (*adf*): (parametric)	-2.198**	-2.137**
Group ρ – Statistic	-1.971**	-2.654**
Group t – Statistic: (non–parametric)	-0.619	-1.295
Group t – Statistic (*adf*): (parametric)	-3.172*	-3.587*

Note*:* – All statistics are from Pedroni's procedure (1999) where the adjusted values can be compared to the N(0,1) distribution. The Pedroni (2004) statistics are one-sided tests with a critical value of –1.64 ($k < -1.64$ implies rejection of the null), except the v – statistic that has a critical value of 1.64 ($k > 1.64$ suggests rejection of the null).

– * and ** indicate rejection of the null hypothesis of no cointegration at one percent and five percent levels of significance, respectively.

Once we have confirmed that there is a presence of a long run relationship between *FDI* inflows into China and independent variables, we have done coefficients estimation. The results of estimation are shown in Table 4.

Table 4: FMOLS Regression Results

FDI Inflows from	Variables			
	OPENNESS	*RELEXC*	*FCF*	*GDP*
Indonesia	5.55* (3.86)	–0.42 (–0.60)	–22.72* (–3.82)	30.85* (4.17)
Malaysia	6.33* (2.54)	0.52 (0.30)	2.91** (2.17)	1.52 (1.01)
Philippines	1.58 (1.03)	0.49 (0.27)	1.30 (0.45)	2.22 (0.89)
Singapore	–4.67* (–3.22)	–0.66 (–1.10)	0.37 (0.45)	3.97* (4.95)
Thailand	0.65 (0.64)	–0.46 (–0.34)	–2.89** (–2.09)	5.74** (2.59)
Panel Group	1.89** (2.17)	–0.11 (–0.66)	–4.21 (–1.27)	8.86* (6.08)

Note: – The null hypothesis for the t – ratio is $H_0 : \beta_i = 0$.
– Figures in parentheses are t – statistics.
– * and ** indicate rejection of the null hypothesis at one percent and five percent levels of significance, respectively.

As shown by Table 4, the sign of estimated coefficient for *OPENNESS* are correct and statistically significant at one percent level for Indonesia and Malaysia, respectively. One unit increase in trade-intensity ratio caused as much as USD5.5 and USD6.63 of *FDI* inflows into China from Indonesia and Malaysia. Higher degree of openness economy of China means the restriction to trade is low or falling, thus facilitating more international exchange of capital between China and these two countries. For the case of the Philippines and Thailand, even though the estimated coefficients show positive relationships with the value of coefficients are 1.58 and 0.65, respectively but they are not significant. In contrast, *OPENNESS* has negative effect on *FDI* inflows into China from Singapore since the estimated coefficient is negative and statistically significant. It means that a higher degree of openness implies a lower level of *FDI* inflows. This can be explained by looking at the type of investments made by the countries. If it is horizontal (market-seeking) investments, trade restrictions (means less openness) would have a positive effect on *FDI* inflows. This is due to the 'tariff jumping' hypothesis which argues that foreign firms that seek to serve local markets may decide to set up subsidiaries in the host market as it is difficult to import their products to the country. The other way around multinational firms engaging in export-oriented investments will prefer to locate in a more open economy since increased imperfections that accompany trade protection imply higher transaction costs associated with exporting.

Furthermore, all coefficients of *RELEXC* are not statistically significant. Therefore, there is no statistically significant long–run relationship between *RELEXC* and *FDI* inflows into China from ASEAN–5. This finding suggests that the changes in relative exchange rates do not influence significantly *FDI* inflows into China from ASEAN–5, which is contrast to the explanation provided by Xu and Wang (2007). Our result shows that the movement of exchange rates plays no role in explaining the level of *FDI* inflows into China. This can be justified with production flexibility argument (De Meza & van der Ploeg, 1987; Aizenman, 1992; Phillips *et al*. 2008). According to this argument, foreign investors are risk neutral and they commit to domestic and foreign capacity *ex ante*, after the realization of nominal or real shocks. This result is consistent with Bailey and Tavlas (1991) and Goldberg and Kolstad (1995). Goldberg and Kolstad, for instance, argued that if investors are classified as risk neutral, there is no significant relationship between exchange rate volatility and the allocation of production facilities between local and foreign markets.

However, this result is also contrast to the common argument of economic theory and available literature on exchange rates and *FDI* that based on risk-aversion argument, imperfect capital market hypothesis, and irreversible investment decision (Dixit & Pindyck, 1994). Economic theory states that as a depreciated exchange rate lowers the cost of production and higher competitiveness of country. Therefore, investment activities particularly *FDI* inflows into the host countries will increase. Furthermore, risk-aversion argument and imperfect capital market hypothesis theoretically claim that the level or exchange rate may negatively influence *FDI*. Using imperfect capital market consideration, study of Froot and Stein (1991) provide the evidence of negative relationship between exchange rates and *FDI*. A weaker host country currency, inward *FDI* tends to increase within an imperfect capital market model as depreciation makes host country assets less expensive relative to assets in the home country. Walsh and Yu (2010) have also stated the similar argument. A weaker real exchange rate might be expected to increase vertical *FDI* as firms take advantage of relatively low prices in host markets to purchase facilities or, if production is re-exported, to increase home-country profits on goods sent to a third market. Meanwhile, depends on the expectations of future profitability, Campa (1993) predicted that an appreciation of the host currency will increase *FDI* into the host country. Using data on Japanese acquisitions in the USA, Blonigen (1997) has concluded different argument from Froot and Stein. Other numerical studies by Cushman (1988) and Dewenter (1995) have supported Campa and Blonigen's views.

Interestingly, the positive coefficient of *FCF* and it is statistically significant at five percent level implies that increase USD1 million of *FCF* caused increment of *FDI* inflows into China from Malaysia as much as USD2.91 million. Thus, this finding is in line with the theory that explains countries with high capital formation normally have more access to capital and they tend to become home country and do investment abroad. On the other hand, *FCF* has negative effect on *FDI* inflows from Indonesia and Thailand since its coefficients values are negative and statistically significant at one percent and five percent level. The result supports the argument of Adhikary (2011). He argued that the level of capital formation is likely to influence FDI and economic growth as well. Neo-classical growth model postulates that developing economies that have a lower initial level of capital stock tend to have higher marginal rate of returns (productivity) and growth rates if adequate capital stock is injected

As expected, we found that the coefficients of *GDP* are positive and statistically significant at one percent level only. Therefore, these results suggest that *GDP* of China has positive effect on *FDI* inflows from Indonesia, Singapore, and Thailand. The results show that increase USD1 million China's *GDP* will cause increase *FDI* inflows from Indonesia, Singapore, and Thailand as much as USD30.85 million, USD3.97 million, and USD5.74 million, respectively. We can conclude that the high rate of economic growth in China will lead to increase the inflows of *FDI* inflows from these countries. *FDI* inflows from

Indonesia are the largest contribution among them. The result obtained from this variable is also similar to studies done by Tsai (1994) and Schneider and Frey (1985). Both studies indicate that real *GDP* per capita as a proxy of market size has a strong positive relationship with *FDI*. However, it should be noted that while market size is still a significant variable, its importance has been slowly decreasing recently. This result is also echoed by (UNCTAD 1996; 97) which found that *FDI* in developing countries are slowly shifting from resource and market seeking to more (vertical) efficiency seeking.

5. Policy Implication and Conclusion

This study examines the determinants of *FDI* inflows into China from ASEAN–5 using the panel cointegration approach. The unit root test (IPS) is used to confirm the stationarity of all variables before performing cointegration test. After confirming that all variables are non–stationary at level, the panel cointegration approach is applied. Using Pedroni's approach, the long–run cointegration test is performed to investigate the existence of the long–run cointegration among the variables. This paper provides statistical evidence that the long–run determinants of *FDI* inflows into China may differ among ASEAN–5 countries. Results obtained indicate the presence of the long–run relationship between *OPENNESS* and *FDI* inflows from Indonesia, Malaysia and Singapore. However, there is no statistically significant long–run cointegration between *RELEXC* and *FDI* inflows from ASEAN–5 countries. For the case of Malaysia, the coefficient of *FCF* is statistically significant with the value of 2.91 and therefore, we can conclude that there is still a long run cointegration between *FCF* and *FDI*.

The results suggest that the macroeconomic policy framework is important. China can indeed usefully undertake policies to foster the volume of *FDI* inflow from the ASEAN–5. In specific, we found that *GDP* and *FCF* are crucial as part of future policy to further attract new *FDI* to inflow to China. The low impact of *GDP* and *FCF* does suggest that China is yet to embark seriously in this issue. This is because, these independent variables could also serve as a cost of doing business and improvement would surely be able to reverse the inflows into China. There is room for China to improve its exchange rate, *GDP*, and fixed capital formation as attraction for *FDI* to inflows into China. This could be explained by the fact that although the ASEAN as a whole is not as populous as China, the purchasing power of the ASEAN is actually higher than China and therefore, could serve as a good avenue for high quality product of multinational corporations. Finally, the impact of trade openness is positive and significant, implying that maintaining the level of openness could be another good policy.

Due to positive development after accession of China in WTO, the ASEAN–5 countries need to explore the possible strategies for accelerating the flow of *FDI* to China. Since *FDI* is viewed as one of the major stimuli of cooperation, some areas to the cooperation among them should be identified and activated such as economic and investment cooperation, trade partnership, and technology development exchanges. Economic and investment cooperation, for instance, will contribute excellent prospects and better opportunities in a wide range of fields for promoting interaction, tightening relationship, and achieving mutual benefits in trade, investment, and industrial and commercial fields between China and ASEAN–5 countries as well as protecting the interests of all countries. The implementation of the ASEAN-China Free Trade Agreement in 2010, for example, has undoubtedly boosted bilateral trade and investment relations between China and ASEAN–5 countries.

Although China's accession into WTO will have an important impact on the ASEAN countries, especially when China's *FDI* intake and global share have increased in a period when global *FDI* inflows have contracted, China's rise should be interpreted as twin opportunity. This is because while *FDI* further promotes China's export, it also expands its domestic market where the ASEAN countries can directly benefit by recording surging exports to China. At the same time, China's cost competitiveness forces the ASEAN countries to move up the value chain by shifting from low technology manufacturing to higher value added, pushing for economics as well as industrial transformation. It would take several years before China exports *FDI* in significant amounts. The ASEAN countries should take this opportunity to actively pursue, instead of encumber economic cooperation and free trade with China.

In short, the empirical results confirm that inflows of *FDI* into China by the ASEAN–5 countries are significantly influenced by all selected variables except exchange rate. Therefore, the macroeconomic fundamental changes are important factors that attract the ASEAN–5 to invest capital in China. The ASEAN–5 countries have also an important stake in encouraging *FDI* inflows in China. For sustaining and enhancing good cooperation between China and the ASEAN–5 countries, the future flow of *FDI* from these countries need to be maintained through mutual benefit agreement and effective foreign investment policies.

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