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## THE ENDOGENEITY OF THE OPTIMUM CURRENCY AREA CRITERIA IN EAST ASIA

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### ABSTRACT

The Asian financial crisis in mid-1997 has increased interest in policies to achieve greater regional exchange rate stability in East Asia. It has renewed calls for greater monetary and exchange rate cooperation. A country's suitability to join a monetary union depends, inter alia, on the trade intensity and the business cycle synchronization with other potential members of the monetary union. However, these two Optimum Currency Area criteria are endogenous. Theoretically, the effect of increased trade integration (after the elimination of exchange fluctuations among the countries in the region) on the business cycle synchronization is ambiguous. Reduction in trade barriers can potentially increase industrial specialization by country and therefore resulting in more asymmetry business cycles from industry-specific shocks. On the other hand, increased trade integration may result in more highly correlated business cycles due to common demand shocks or intra-industry trade. If the second hypothesis is empirically verified, policy makers have little to worry about the region being unsynchronized in their business cycles as the business cycles will become more synchronized after the monetary union is formed. This paper assesses the dynamic relationships between trade, finance, specialization and business cycle synchronization for East Asian economies using a Generalized Method of Moments (GMM) approach. The dynamic panel approach improves on previous efforts to examine the business cycle correlation – trade link using panel procedures, which control for the potential endogeneity of all explanatory variables. Based on the findings on how trade, finance and sectoral specialization have effects on the size of common shocks among countries, potential policies that can help East Asian countries move close toward a regional currency arrangement can be suggested. The empirical results of this study suggest that there exists scope for East Asia to form a monetary union.

*Keywords:* Optimum Currency Area; Monetary Union; Trade Integration; Business Cycle Synchronisation

*JEL codes:* E3, F1

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## 1. INTRODUCTION

While a managed float exchange rate regime is the trend in the modern international monetary system, there is also a movement towards the other extreme of a credible bilateral or multilateral parity-fixing regime. In particular, formation of the European Economic and Monetary Union (EMU) with twelve countries adopting a single currency (Euro) has sparked an intense debate on monetary union. Interest on a monetary union has spilled over to other regions in the world including East Asia. Ever since the Asian Financial Crisis in 1997, East Asian countries have had a strong impetus to search for a regional exchange rate coordinating mechanism that could help secure financial stability in the region. This search has opened the door to the possibility of a common currency and exchange rate system to promote greater economic integration and monetary cooperation in the region. If East Asian countries were interested to pursue the idea of a common regional currency, detailed studies on this issue must be carried out. Robert Mundell's theory of Optimum Currency Area (OCA) has been widely used in literature to assess the suitability of a common currency. The criteria include the symmetry of shocks across countries, factor mobility, wage flexibility, trade and financial integration, and political integration. The greater the linkages between the countries using any of the above criteria, the more suitable a common currency.

However, the first two OCA criteria are endogenous. A region's business cycle may become more synchronised after a monetary union is formed. This can be explained by the increased intra-regional trade (after the elimination of exchange fluctuations among the countries in the region) that potentially increases the business cycle synchronization. If such hypothesis is empirically verified, policy makers have little to worry about the region being unsynchronised in their business cycles as the business cycles will become more synchronised after the monetary union is formed. Therefore, it is imperative to investigate the relationship between business cycle synchronization and trade integration. Most of the existing literatures investigate the relationship between trade and business cycle using simple instrumental variables method with time series data. A recent study by Imbs (2004) has improved on the methodology by using a simultaneous approach. However, there are still estimation problems to the model as the variables are very likely to be endogenous. In addition, heteroskedasticity will add problems to the estimation. We proposed a GMM approach to estimate the model as it will solve the problems encountered in the previous studies. In addition, our estimation using panel procedures will improve on previous efforts to examine the business cycle correlation – trade link the following ways: (1) estimation using panel data allows us to exploit the time-series and cross section nature of the dynamic relationships between trade, finance, specialization and business cycle synchronization; (2) in a pure cross-country instrumental variable regression, any unobserved country-specific effect becomes part of the error term, which may bias the coefficient estimates; (3) unlike most existing cross-country studies, our panel estimator controls for the potential endogeneity of all explanatory variables. Based on the findings on how trade, finance and sectoral specialization have effects on the size of common shocks among countries, potential

policies that can help East Asian countries move close toward a regional currency arrangement can be suggested.

## 2. METHODOLOGY

We investigate the following equation:

$$Y(s)_{ij,t} = \alpha_0 + \alpha_1 Y(S)_{ij,t-1} + \alpha_2 T(\omega)_{ij,t} + \alpha_3 S_{ij,t} + \alpha_4 F_{ij,t} + \varepsilon_{ij,t} \quad (1)$$

$Y(s)_{ij,t}$ : denotes the bilateral business cycle correlation between country  $i$  and  $j$ , *de-trended with method  $s$*  (corresponding to Band-Pass filtering and HP-filtering)

$T(\omega)_{ij,t}$ : denotes the bilateral trade intensity between country  $i$  and  $j$  using trade intensity concept  $\omega$  (corresponding to: total bilateral trade normalised by GDP or a measure based on Deardorff (1998))

$S_{ij,t}$ : a specialization index capturing how different the sectoral allocations of resources are between country  $i$  and  $j$

$F_{ij,t}$ : a measure of financial integration for each country pairs

For equation (1) we assume that the  $\varepsilon_{it}$  follow a one-way error component model

$$\varepsilon_{it} = \eta_i + v_{it}$$

where  $\eta_i \sim \text{i.i.d. } (0, \sigma_\eta^2)$  and  $v_{it} \sim \text{i.i.d. } (0, \sigma_v^2)$ , independent of each other and among themselves.

## 3. THE MEASUREMENT

Bilateral correlations in business cycles are computed based on the basis of the cyclical component of real economic activity. While most of the researchers use real GDP as a measure of real economic activity, an index of industrial production, total employment, and the unemployment rate are sometimes used as measures of real economic activity.<sup>2</sup> Although it is tempted to employ all the four different measures of real economic activity, data restriction constrains us to employ only the real GDP. The real GDP will then be transformed in two different ways. Firstly, it will be transformed to the natural logarithm. Secondly, the variable will be de-trended so as to focus on business cycle fluctuations. We employ the commonly used Band-Pass (BP) Filter introduced in the Hodrick-Prescott (HP) filter.<sup>3</sup>

Two different measures of trade will be considered. The first one is a standard measure used in many recent studies, for instance Clark and van Wincoop (2001),

<sup>2</sup> Among others, Frankel and Rose (1997, 1998) use real GDP, an index of industry production, total employment and unemployment rate as measures of real economic activity.

<sup>3</sup> This is a widely used de-trending technique that is also employed by Choe (2001), Frankel and Rose (1997, 1998), and Imbs (2004) among others. We apply the traditional smoothing parameter of 100 for our annual data.

Frankel and Rose (1997, 1998), and Imbs (2004) among others. This standard measure is used for benchmarking purpose. The bilateral trade intensity is calculated as:<sup>4</sup>

$$T_{i,j}^1 = \frac{1}{T} \sum_t \frac{X_{i,j,t} + M_{i,j,t}}{Y_{i,t} + Y_{j,t}} \quad (2)$$

where  $X_{i,j,t}$  denotes total nominal exports from country  $i$  to  $j$  during period  $t$ ,  $M_{i,j,t}$  represents imports to  $j$  from  $i$ , and  $Y_{i,t}$  denotes the level of nominal GDP in country  $i$  at period  $t$ . The second measure is suggested by gravity models, also employed by Clark and van Wincoop (2001) and Imbs (2004) among others.<sup>5</sup> In this measure, bilateral trade between country  $i$  and  $j$  is multiplied by a scale factor of world GDP and then divided by the product of the GDPs of  $i$  and  $j$ :<sup>6</sup>

$$T_{i,j}^2 = \frac{1}{2} \frac{1}{T} \sum_t \frac{X_{i,j,t} + M_{i,j,t}}{Y_{i,t} \times Y_{j,t}} (Y_t^W) \quad (3)$$

Where  $Y_t^W$  is world GDP during period  $t$ .  $T^2$  differs from  $T^1$  in that it is independent of country size and depends only on trade barriers.<sup>7</sup> In practice, we take the natural logarithm of both ratios.

There is no standard measure of similarity in industry specialization. Imbs (2001) uses a correlation coefficient between sectoral shares in aggregate output or employment, whereas Imbs (2004) use sectoral real value added to compute the industry specialization. We adopt the Herfindahl index of concentration adopted in Krugman (1991) and Clark and van Wincoop (2001). The similarity of country  $i$  and  $j$ 's production structures is measured as:

$$S_{i,j} = \frac{1}{T} \sum_t \sum_n^N |S_{ni} - S_{nj}| \quad (4)$$

where  $S_{ni}$  and  $S_{nj}$  denote the GDP shares for industry  $n$  in country  $i$  and  $j$ . If the two countries had identical industrial structures, that is, that industry shares of GDP were the same for country  $i$  and  $j$ , then the index would be zero. The index reaches its maximal value of two when two countries have no sector in common (because each share in each region would be counted in full). Therefore the index is a rough way of quantifying differences in structures and specialization between countries.  $S_{i,j}$  is measured in natural logarithm in practice.

There are a few measures of financial integration. However, due to data restriction, we follow Imbs' (2004) measures of financial integration through four binary variables reporting the number of countries with: (i) multiple exchange rates, (ii) currency account restrictions, (iii) capital account restrictions and (iv) surrender of export proceeds as reported in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). The composite index from AREAER is averaged

<sup>4</sup> Frankel and Rose (1997, 1998) also use another trade measure that is normalized by total trade between country  $i$  and  $j$ .

<sup>5</sup> This measure is based on Deardorff (1998).

<sup>6</sup> They drop the scale factor of world GDP in Deardorff's (1998) model because they take the logarithm of the trade variable (and the scale factor will be absorbed in the constant of the regression).

<sup>7</sup> Deardorff shows that  $T^2$  equals to 1 if there are no trade barriers and if preferences are homothetic.

each year, and thus can take values 0, 0.25, 0.5, 0.75 or 1. It is then summed for each country pairs before taking the time averages as the other variables discussed above.

We compute bilateral correlations for the cyclical components of real activity. These correlations are estimated over a given span of time. We follow the standard procedure of splitting our sample into five parts: 1970 – 1976; 1977 – 1984; 1985 – 1992; 1993 – 1999; 2000-2006.<sup>8</sup>

This study examines 8 East Asian countries, namely ASEAN 5 – Indonesia, Malaysia, the Philippines, Singapore and Thailand;<sup>9</sup> plus three additional members of East Asia – China, Japan and Korea. Although the intention is to study all the East Asian countries, unfortunately we lack data on Brunei, Cambodia, Laos, Myanmar and Vietnam. *GAUSS* program is used in the estimation.

The bilateral trade data are from the IMF's Direction of Trade Statistics. We use data on real GDP and gross value added classified at 7 broad categories (ISIC one digit) from United Nation's National Accounts Main Aggregates Database to compare the differences in the sectoral composition of GDP between an economy and the rest of economies in the region.<sup>10</sup> The 7 broad industries are agriculture, mining, manufacturing, construction; wholesale, transport and other activities. Output is measured by the log of real GDP growth.

Financial integration is measured through four binary variables reporting the number of countries with: (i) multiple exchange rates, (ii) current account restrictions, (iii) capital account restrictions and (iv) surrender of export proceeds as reported in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER).

#### 4. POTENTIAL PROBLEMS OF ESTIMATION

The proposed linear regression equation (1) poses some challenges for estimation for a few reasons. Firstly, it is common for panel data sets, consisting of cross sections observed at several points in time to exhibit both characteristics of heteroscedasticity and autocorrelation.<sup>11</sup> Since we are looking at different countries with heterogeneous characteristics, the variances of the regression disturbances are not likely to be constant across observations. Secondly, the explanatory variables are potentially endogenous and the use of lagged  $Y$  on the right hand side induces correlation between the regressor and the error in first differences. Since  $Y_{it}$  is a function of  $\eta_i$ , it immediately follows that  $Y_{i,t-1}$

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<sup>8</sup> This time period is chosen because the sectoral data from United Nation is only available from 1970.

<sup>9</sup> The new ASEAN members include Cambodia, Laos, Myanmar and Vietnam are excluded in the study as the stages of development in these countries are very much different from the rest of the ASEAN members. Williamson (1999), for example, omits the new members of ASEAN, limiting the heterogeneity of the countries adopting a common basket peg.

<sup>10</sup> The National Accounts Main Aggregates Database is available from the website: <http://unstats.un.org/unsd/snaama/dnllist.asp>

<sup>11</sup> Imbs' (2004) simultaneity methodology may suffer from this problem.

is also a function of  $\eta_i$ . Therefore,  $Y_{i,t-1}$  is correlated with the error term. This implies that the OLS estimator is biased and inconsistent even if  $v_{it}$  are not serially correlated.

## 5. THE GENERALISED METHOD OF MOMENTS (GMM) ESTIMATION

The use of GMM estimators suggested by Arellano and Bond (1991) is one common method used to control for the above mentioned problems arising in dynamic panel data models. The main idea behind this method is that additional instruments can be obtained if one utilizes the orthogonality conditions that exist between lagged values of  $Y_i$  and  $v_{it}$ . This approach has the advantage of avoiding biases related to omitted specific individual effects and to control for endogeneity. In addition, GMM estimator does not require any particular distributions of the error term. Hence, even in the presence of heteroscedasticity, the estimator produces consistent and efficient estimates of the unknown parameters.

### 5.1 *Nonstationary Dynamic Panel and Panel Unit Root Tests*

So far, we have not addressed an important estimation issue that parameters may not be identified using first-differenced GMM estimators when the series are random walks, and more generally identification may be weak when the series are near unit root processes. As shown in Blundell and Bond (1998), this can result in large finite-sample biases when using the standard first-differenced GMM estimator. It is recognised that the GMM estimator using alone the moment restrictions suffers a downward bias when  $\alpha$  is near to unity. Binder *et al.* (2000) also show that the conventional GMM estimators based on standard orthogonality conditions break down if the underlying time series contain unit roots. Therefore, it is imperative to perform unit root test on our series. Blundell and Bond (1998) revealed that the system GMM estimator using additional moment restrictions ameliorates the downward bias, in their theoretical illustrations and Monte Carlo experiments. This section provides the unit roots identification in our panel data models. In addition, the system GMM estimator that is suggested to cope with the problem of nonstationarity will be discussed.

We employ two commonly used panel unit root tests: (i) Levin *et al.* (2002), hereafter denoted as LLC and (ii) Hadri (2000). LLC tests the null hypothesis of non-stationarity while Hadri tests the null hypothesis of stationarity. In this section we briefly discuss the panel data unit root tests of LLC and Hadri respectively.

## 5.2 System GMM

Let us rewrite equation (1) as:

$$Y(s)_{ij,t} = \alpha_0 + \alpha_1 Y(S)_{ij,t-1} + \beta X_{i,t} + \eta_i + \nu_{i,t} \quad (5)$$

$X_{i,t}$  is a vector of explanatory variables, that is  $X = [T(\omega), S \text{ and } F]$ .  $\eta_i$  is an unobserved represents country specific effect which can be eliminated by first differencing equation (5):

$$\Delta Y(s)_{ij,t} = \alpha_1 \Delta Y(S)_{ij,t-1} + \beta \Delta X_{i,t} + \Delta \nu_{i,t} \quad (6)$$

Blundell and Bond (1998) show that the biases caused by near unit root processes can be dramatically reduced by exploiting reasonable stationarity restriction on the initial condition processes. The system GMM estimation combines the standard set of equations in first-differences with suitably lagged levels as instruments, with an additional set of equations in levels with suitably lagged first-differences as instruments. Both sets of moment conditions can be exploited as a linear GMM estimator in a system containing both first-differenced and levels equations. Combining both sets of moment conditions provides what we label the system GMM estimator. We adopted this system GMM estimator to provide a consistent estimation of our model. Based on Blundell and Bond (1998) framework, we used both the levels and differences of the lagged  $Y(s)$ ,  $T(\omega)$ ,  $S$  and  $F$  as instrumental variables. As an empirical matter, the validity of these additional instruments can be tested using standard Sargan tests of over-identifying restrictions, or using Difference Sargan or Hausman comparisons between the first-differenced GMM and system GMM results (see Arellano and Bond, 1991).

## 6. RESULTS AND ESTIMATION

This section presents the empirical results and discussions of the examination of the new OCA theory using the GMM approach. We first test for the presence of random walks by conducting the panel unit root tests. Then we test for the validity of the system GMM instruments employed in the model. Finally, we report the system GMM estimation.

### 6.1 Nonstationary Dynamic Panel and Panel Unit Root Tests

This section provides the common panel root tests on the series as a standard procedure. Such tests are important as the presence of random walks in the series will cause downward biasness in the estimation using the standard GMM estimator. We employ two commonly used panel unit root tests: (i) Levin *et al.* (2002) and (ii) Hadri (2000). LLC tests the null hypothesis of non-stationarity while Hadri tests the null hypothesis of stationarity. Table 1 reports these tests results. LLC is commonly used panel unit root tests. However, it suffers from the lack of power (Hadri 2000) since the

null hypothesis of a unit root tends to be accepted unless there is strong evidence to the alternative. This is one form of type II error (Davidson and MacKinnon 1993; Greene 2003). It is therefore recommended to test a null of unit root as well as a null of no unit root. We therefore employ a well-known test for the null of no unit root that is proposed by Hadri (2000).

Table 1 presents the results of the panel unit root tests. Under the LLC, the null hypothesis of non-stationarity for all the level series could not be rejected, but the null hypothesis of non-stationarity for all the first differenced series is rejected. For instance, under the LLC test, the level  $Y(BP)$  without trend has a  $p$ -value of 0.92, implying that the hypothesis of non-stationarity could not be rejected. However, after first differencing, the null hypothesis of non-stationarity is rejected (i.e., LLC test for  $Y(BP)$  after first differencing has a  $p$ -value of 0, implying the rejection of the null of non-stationarity at 1 per cent significance level). By employing the Hadri (2000) test, whether or not trend is included in the test equation gives different results. For series  $T(2)$  and  $S$  in particular, the null hypothesis of stationarity cannot be rejected even at the level equation with no trend is included. After plotting the series, we did find trend in the series. Therefore, if we base our conclusion on the test results with trend for the above-mentioned series, we reject the null hypothesis of no unit root under the levels but we could not reject the null of stationarity under the first differences for all series. Both LLC and Hadri tests have consistently implied that all the series are  $I(1)$  variables.



**Table 1 The Panel Unit Root Tests**

Variable	Level				First Difference		
	LLC		Hadri		LLC	Hadri	
	No Trend	Trend	No Trend	Trend	None*	No Trend	Trend
Y(HP)	0.65 [1] (0.74)		1.93** [1] (0.03)		-10.51*** [1] (0.00)	-1.08 [1] (0.86)	
T(1)	0.56 [1] (0.71)	1.24 [1] (0.89)	9.68*** [1] (0.00)	5.26*** [1] (0.00)	-10.83*** [1] (0.00)	-1.06 [1] (0.86)	1.40 [1] (0.92)
T(2)	-0.89 [1] (0.19)	-0.31 [1] (0.38)	0.26 [1] (0.40)	1.47* [0] (0.07)	-8.28*** [1] (0.00)	-1.04 [1] (0.85)	1.32 [1] (0.91)
S	-0.64 [1] (0.26)	-0.11 [1] (0.46)	0.12 [1] (0.45)	1.94** [1] (0.03)	-9.38*** [1] (0.00)	-1.01 [1] (0.84)	1.21 [1] (0.89)
F	-0.45 [1] (0.33)	-0.03 [1] (0.49)	5.88*** [1] (0.00)	3.51*** [1] (0.00)	-9.57*** [1] (0.00)	-1.02 [1] (0.85)	1.27 [1] (0.90)

Panel unit root tests are based on Levin *et al.* (2002) and Hadri (2000). The null hypothesis of LLC is the presence of unit root, while that of Hadri is no unit root. Figures in parentheses ( ) indicate *p*-values. Figures in [ ] indicate the lag length for LLC test and bandwidth for the Hadri test. Asterisks \*\*\*, \*\* and \* indicate significance at 1%, 5% and 10% level respectively. Note that trend is not included in the test equation for series Y (HP) as these series are de-trended. The bandwidth selected for all series under the LLC is one (using the Bartlett kernel). The bandwidth selection under the Hadri test is using quadratic spectral kernel. Y (HP): bilateral business cycle correlations based on real GDP filtered by HP filter. T (1): bilateral trade normalized by GDP. T (2) bilateral trade measure by Deardorff (1998).

\* No trend no intercept

## 6.2 Test for the Validity of the System GMM Instruments

Since all our series are  $I(1)$ , the first-differenced GMM estimator will not be efficient. The lagged levels are used as instruments for the equations in the first-differences. Since all our series are non-stationary in levels, lagged levels of the series are not suitable instruments in this context. The level instruments for the first differenced equations will tend to be weak as the lagged levels are weakly correlated to subsequent first-differences, the consequence of which is serious finite sample biases (Blundell and Bond 1998, 2000).

In view of the shortcomings associated with the first-differenced GMM estimator, the system GMM estimator is being employed in our study. In addition to using lagged levels in the equations for first differences, the lagged differences of variables are also used as instruments in equation for levels. However, there is a risk of over-identifying restrictions as a range of moment conditions are identified. The notable test for the validity of the instruments for system GMM is the difference Sargan test, which is  $\chi^2$  distributed, and under the null hypothesis of valid instruments.

We tested the validity of the instrumental variables (and therefore the validity of the endogeneity assumption of the explanatory variables) using the Sargan difference tests. There are four different sets of results as there are two different measures of bilateral real GDP correlations and trade intensities. The bilateral business cycle correlation between country  $i$  and  $j$  are de-trended using the HP-filtering. The bilateral trade intensity between country  $i$  and  $j$  is measured using a total bilateral trade normalized by GDP and a measure based on Deardorff (1998). The system GMM estimation therefore yields four different sets of results as in Table 2 using these different measures. It is evident that all four sets of estimation pass the Sargan's test for validity of instrumental variables at 5 per cent significance level. However, it should be noted that the  $p$ -value of the Sargan's test is higher (implying greater confidence of acceptance of the null hypothesis) when the HP-filter is used to de-trend the dependent variable  $Y$ . For instance, the  $p$ -values of the Sargan's test for the model which employs the HP-filtered bilateral business cycle correlation with  $T^1$  and  $T^2$  trade measure are 0.34 and 0.45 respectively (referring to *method 1* and *method 2*), implying that the null hypothesis of valid instruments cannot be rejected.

## 6.3 System GMM Estimation

After a currency union is formed, it is likely to increase the intra-regional trade due to the elimination of exchange fluctuations among the countries in the region. Frankel and Rose (1997, 1998) found a strong and robust positive relationship between trade and business cycles synchronization. Their results imply that the examination of historical data may give a misleading picture of a country's eligibility to join a monetary union as the business cycle synchronization is likely to change dramatically as a result of the formation of monetary union. In other words, countries that do not meet the OCA criteria may still join a monetary union as they are likely to meet the criteria only after joining one. However, Frankel and Rose's view on the endogeneity of OCA criteria is not universally accepted. Theoretically, increased

international trade could result in either more or less correlated business cycle. We employ the system GMM estimation to test the endogeneity of OCA hypothesis. In particular, we empirically analyse how the business cycle in East Asia is influenced by trade, sectoral specialization and financial integration in the region.

First, let us examine how increased trade affects the business cycles in the region. The object of interest to us is the sign and the size of the slope coefficient  $\alpha_2$  in equation (1). The sign of the slope tells us whether the ‘specialization’ effect dominates (in which case we would expect a negative  $\alpha_2$ ) or the ‘endogeneity hypothesis’ dominates (in which case we would expect a positive  $\alpha_2$ ). The size of the coefficient then allows us to quantify the economic importance of this effect. The system GMM estimation of equation (1) is shown in Table 2. The estimated coefficients, along with their standard errors and the level of significance are presented, corresponding to the two different measures of bilateral business cycle correlations and the two different measures of bilateral trade intensity. There are therefore two different sets of estimates which we label *method 1*, and 2. The coefficients when bilateral trade intensity is normalized by the product of GDPs are lower because the scale of the variable is much different, as the ratio of trade to the product of GDPs varies widely by country and time.<sup>12</sup>

The sign of the trade coefficient  $\alpha_2$  in equation (1) is estimated to be positive and statistically significant for both trade measures. The sizes of the coefficients are 0.001 and 0.0001 respectively. Kose and Yi (2002) have calibrated and simulated a three-country business cycles model with transportation costs and technology shocks. Their model yields simulated values for  $\alpha_2$  ranges from 0.0007 to 0.036. On the basis of estimates in Table 2, our estimates of 0.001 appear to be within this range.

The estimates indicate that increased trade increases the business cycle synchronization in the region, implying that the ‘endogeneity effect’ dominates. To give some economic interpretation to the results, we consider the coefficient in *method 1* which is estimated to be 0.001 at 1% level of significance. It means that an increase in  $T_1$  by one standard deviation implies that the bilateral correlation of business cycles would increase by 0.003 from the mean [= (0.001\*2.53)]. The sign of our estimates of  $\alpha_2$  is robust to the different methods of calculating trade intensity that we employ, although the size of the estimates differs according to the methods employed.

Secondly, our empirical results show that the more similar the countries’ economic structure (the lower the  $S$ ) the more correlated the business cycles.  $S$  is a specialization index capturing how different the sectoral allocations of resources are between countries. The estimated coefficients for  $S$  ( $\alpha_3$  in equation 1) shown in Table 2 are highly significant with the expected negative signs, indicating that two economies with a similar economic structure (i.e., lesser degree of specialization) have significantly more correlated business cycles (i.e., lower  $S$  increases the business cycles correlations). The estimates for  $\alpha_3$  using both *method 1* and 2 are -0.017 and -0.010 respectively, it means that an increase in  $S$  by one standard deviation (based on *method 1*), the bilateral correlation of business cycles would reduce by 0.008 [= 0.017\*0.48].

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<sup>12</sup> Frankel and Rose (1998) also reported very different size of parameters (with the same sign) when using different trade measures.

Theoretically, the effects of increased financial integration on business cycle synchronization are ambiguous. On the one hand, increased financial integration allows consumers to cushion against adverse domestic shocks by lending and borrowing abroad. Therefore, volatility of consumption would decline. On the other hand, financial integration increases the potential that domestic financial market distortions get magnified due to foreign capital inflow. Therefore, volatility of output and investment would increase. Both estimates of the finance coefficients  $\alpha_4$  using are 0.016 and 0.009 respectively. The positive sign implies that lower  $F$  (i.e., increased financial integration) leads to lower business cycle correlations. It means that a decrease in  $F$  (indicating increased financial integration) by one standard deviation (based on *method 1*), the bilateral correlation of business cycles would decrease by 0.006 [= 0.016\*0.40]. Business cycles in financially integrated regions are less synchronized according to our estimates. This result is not consistent with that of Imbs (2004) and Otto *et al.* (2001), inter alia, whose studies based on industrial countries found that financial integration positively affects the business cycle synchronization.

Frankel and Rose's (1997, 1998) study on the endogeneity of the OCA criteria has created much interest and debate. A lot of research work has since published to confirm Frankel and Rose's results employing different methodologies and samples. Their pioneered work in this area has become the benchmark of other related studies. Most studies examining the impact of trade on business cycles synchronization find a positive link between the two variables, regardless of the way in which the trade relationship is modeled, confirming Frankel and Rose's conclusion. However, more recent studies tend to find somewhat lower effects than Frankel and Rose (1997, 1998). For instance, using the same sample as Frankel and Rose, Gruben *et al.* (2002) find that the positive trade effect on business cycle correlation is only about half of Frankel and Rose's point estimate. In a recent study using a sample of 21 OECD countries, Inklaar *et al.* (2005) confirm Frankel and Rose's general conclusion but the trade effect on business cycles synchronization is much smaller than reported by Frankel and Rose.

While most of the studies that show a stronger positive link between trade and business cycles synchronization have focused on the developed countries (i.e., Frankel and Rose's (1997, 1998) study is based on twenty one industrialized countries), our study which investigates the East Asian countries has a much weaker positive link. In comparison with Frankel and Rose's estimates, we find much smaller trade effect on business cycles synchronization. The trade coefficient in our study is estimated to be 0.001 compared to Frankel and Rose's estimates of 0.048.<sup>13</sup> Our result is consistent with the findings of Calderon *et al.* (2002) who found that the positive effect that trade has on business cycle synchronization is lesser in the less-developed countries.

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<sup>13</sup> This result is based on the model which employs HP filter with trade normalized by GDP as with Frankel and Rose's model.

**Table 2 System GMM Estimation**

	Trade normalized by GDP ( $T^1$ )	Deardorff's trade measure ( $T^2$ )
	<i>Method 1</i>	<i>Method 2</i>
Trade	0.0010*** (0.0003)	0.0001*** (0.0000)
Specialization	-0.0165*** (0.0023)	-0.0112*** (0.0017)
Finance	0.0155** (0.0017)	0.0093** (0.0009)
<i>p</i> -value for Sargan Test	0.34	0.45

Standard errors in parentheses. Sargan statistics are  $\chi^2$  distributed. We report the *p*-values of the Sargan test for the null hypothesis of valid instruments specification. Asterisks \*\*\* and \*\* indicate significance at 1% and 5% level respectively.

It is generally accepted that the business cycle co-movements are strengthened only when increased trade is accompanied by more intra-industry trade (Shin and Wang 2003 and 2004, Alesina *et al.* 2002 and Imbs 2004). If the increased trade following a currency union is mainly inter-industry trade, the business cycle co-movements can be weakened, therefore making the currency union undesirable *ex post facto* (Shin and Wang 2003). According to Alesina *et al.* (2002), the type of trade between two countries is likely influenced by the levels of per capita GDP and intra-industry trade tends to be much more important for rich countries. Our results which show a weaker positive link between trade and business cycles correlation may be explained by the fact that the intra-industry trade index in East Asia is lower than that of Europe. As reported in Shin and Wang (2003), the intra-industry trade index (IIT) for 3-digit industry classifications in East Asia was only 20.0 compared to the EU's IIT of 46.6.<sup>14</sup> However, the gap becomes much smaller in 1999 with East Asia's IIT of 45.0 versus EU's IIT of 52.3. While the index has monotonically increased in both regions, but the speed is much faster in East Asia. If the hypothesis that the business cycle co-movements are strengthened when increased trade is accompanied by more intra-industry trade is true, and if the tendency of intra-industry trade in East Asia continues to increase, then the business cycle co-movements can be strengthened.

Since increased trade links lead to increased specialization, the positive effect that trade intensity has on business cycle synchronization was mitigated by the negative effect that specialization has on business cycle synchronization.

## 7. CONCLUDING REMARKS

We employ a system GMM approach using panel procedures to examine the endogeneity of OCA criteria. Our model examines the dynamic relationships between trade, finance, specialization and business cycle synchronization. The overall effect of trade on business cycles synchronization is found to be positive, implying that increased trade leads to more synchronized business cycles. This remains true even though increased trade integration results in more specialized economies and less synchronized business cycles as a consequence. Specialization patterns are found to have negative impact on the business cycles synchronization. Since financial integration tends to result in more specialized economies, less synchronized business cycles will be resulted as a consequence.

Although our results show a much weaker positive link between trade and business cycles in East Asia compared with studies which have focused on the developed countries, the fact remains that trade integration increases business cycle synchronization. As such, policy makers may have little to worry about the region being unsynchronized in their business cycles as the business cycles will become more synchronized after the monetary union is formed.

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<sup>14</sup> The East Asian countries included in Shin and Wang's study are China, Hong Kong, Japan, Korea, Taiwan, Indonesia, Malaysia, the Philippines, Singapore, Thailand, Bangladesh and India.

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## APPENDIX A

Diagrams showing the difference between HP-smoothed and unsmoothed real GDP (logarithm difference) data for all countries.

Figure A.1 HP-Filtered Real GDP for China

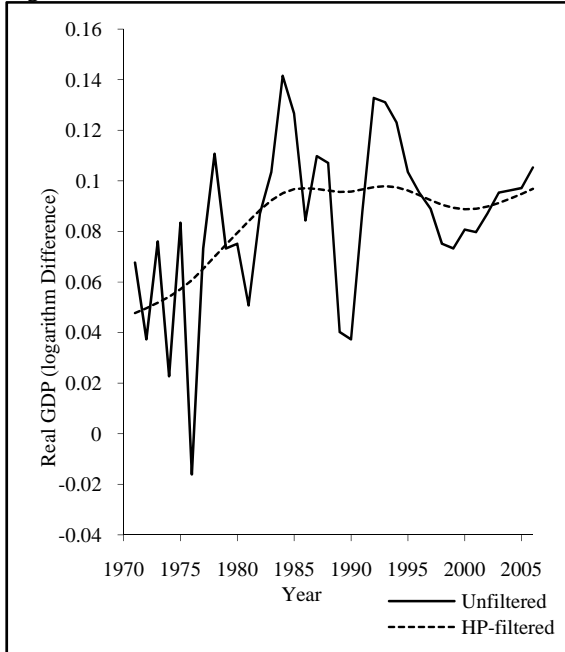


Figure A.2 HP-Filtered Real GDP for Indonesia

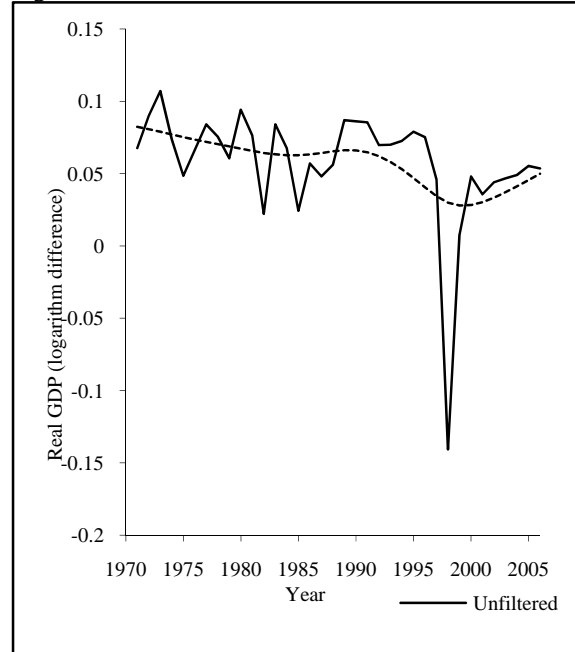


Figure A.3 HP-Filtered Real GDP for Japan

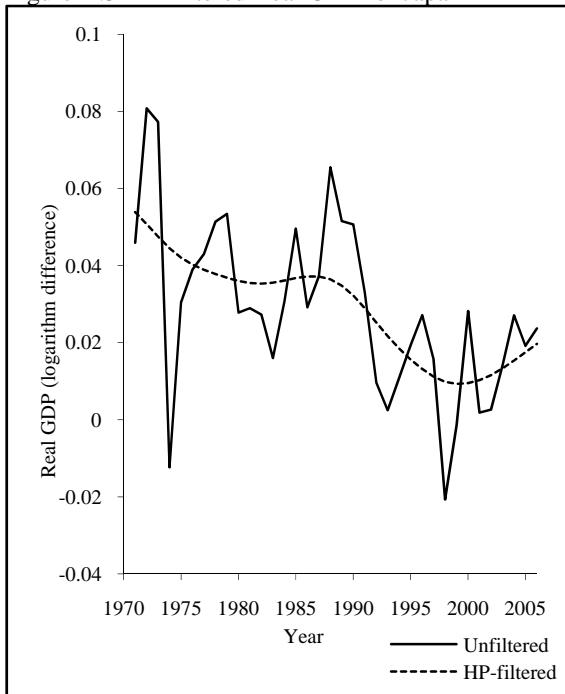


Figure A.4 HP-Filtered Real GDP for Korea

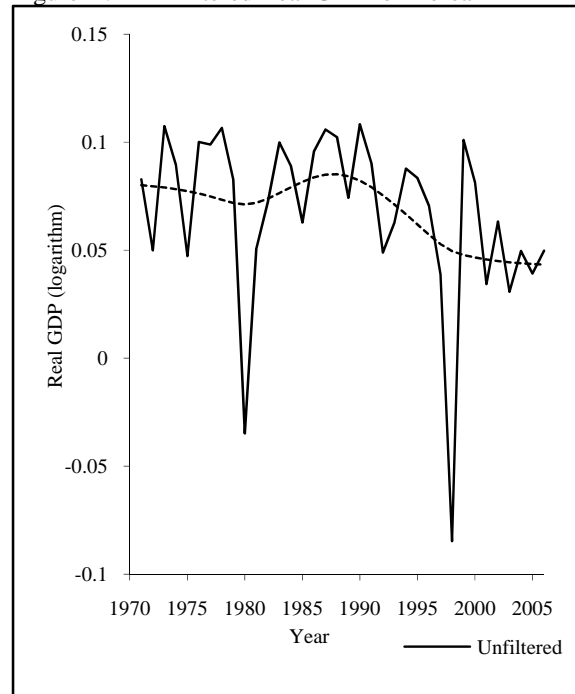


Figure A5 HP-Filtered Real GDP for Malaysia

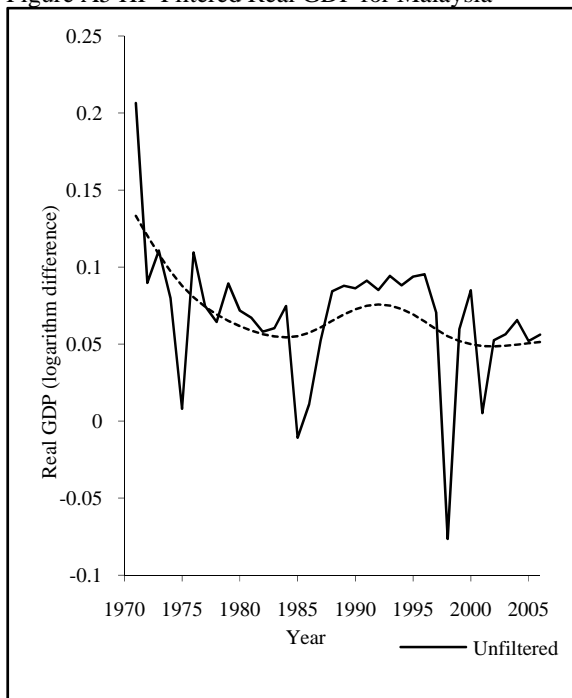


Figure A.6 BP-Filtered Real GDP for Philippines

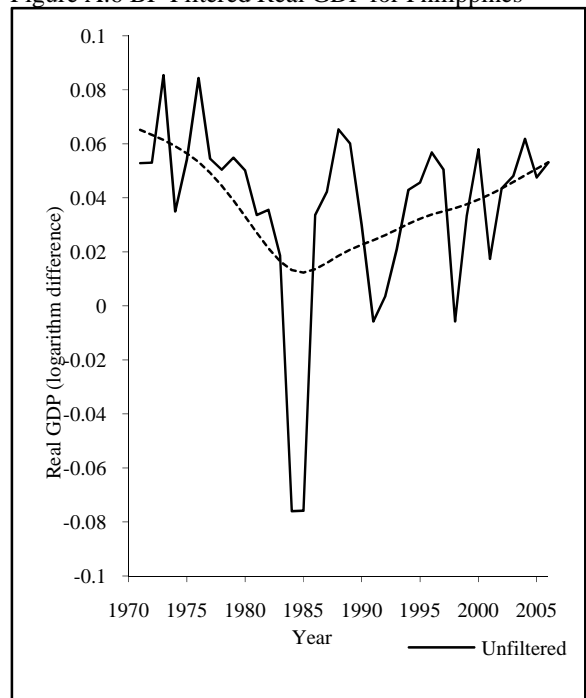


Figure A.7 HP-Filtered Real GDP for Singapore

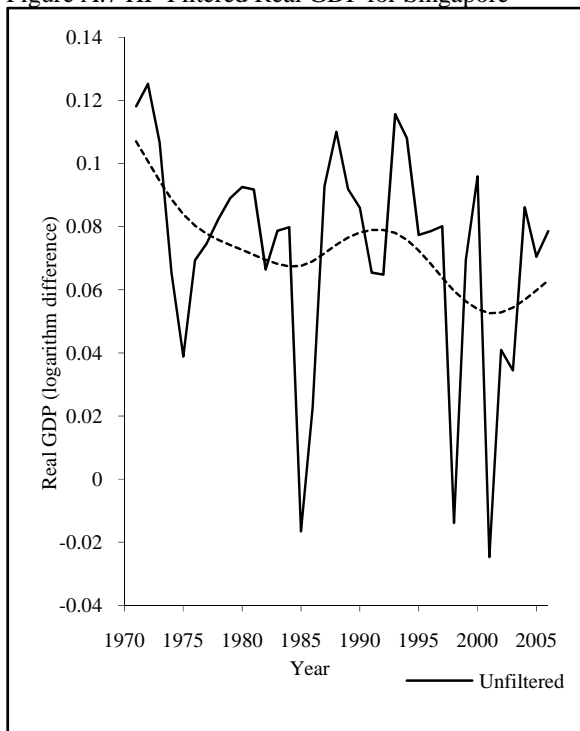
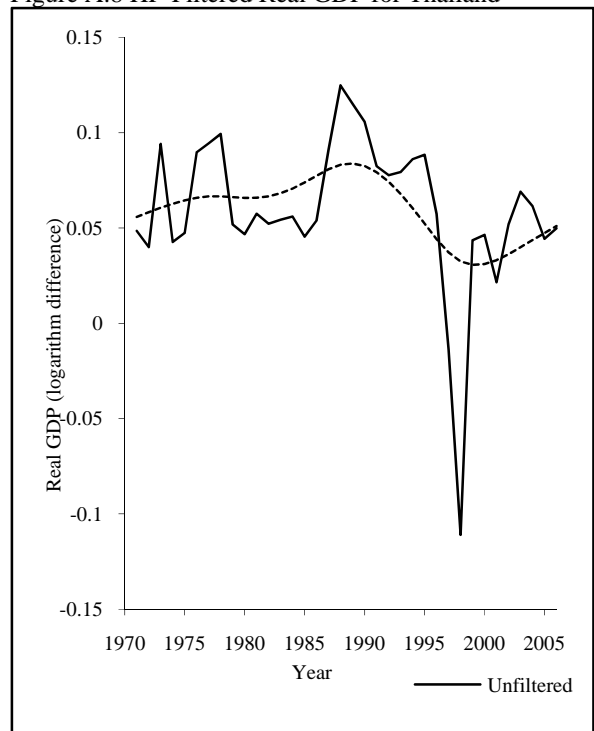


Figure A.8 HP-Filtered Real GDP for Thailand



**APPENDIX B**

Bilateral trade intensity  $T^l$  as calculated in equation (2) for all countries are plotted here.

Figure B.1 Bilateral Trade Intensity  $T^l$  between China and East Asian Countries

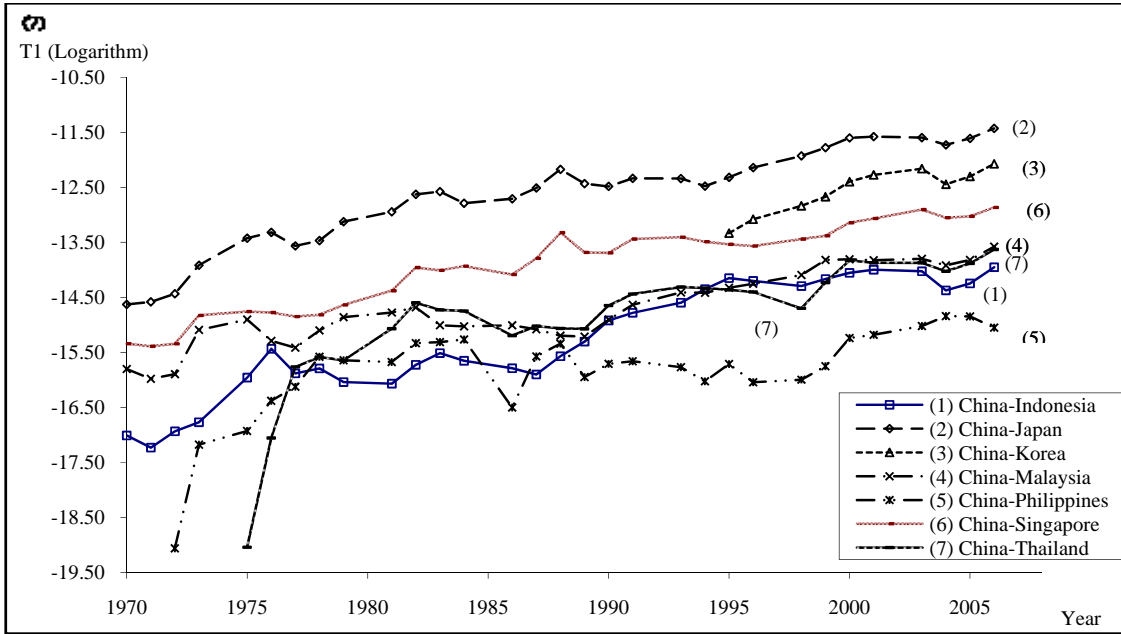


Figure B.2 Bilateral Trade Intensity  $T^l$  between Indonesia and East Asian Countries

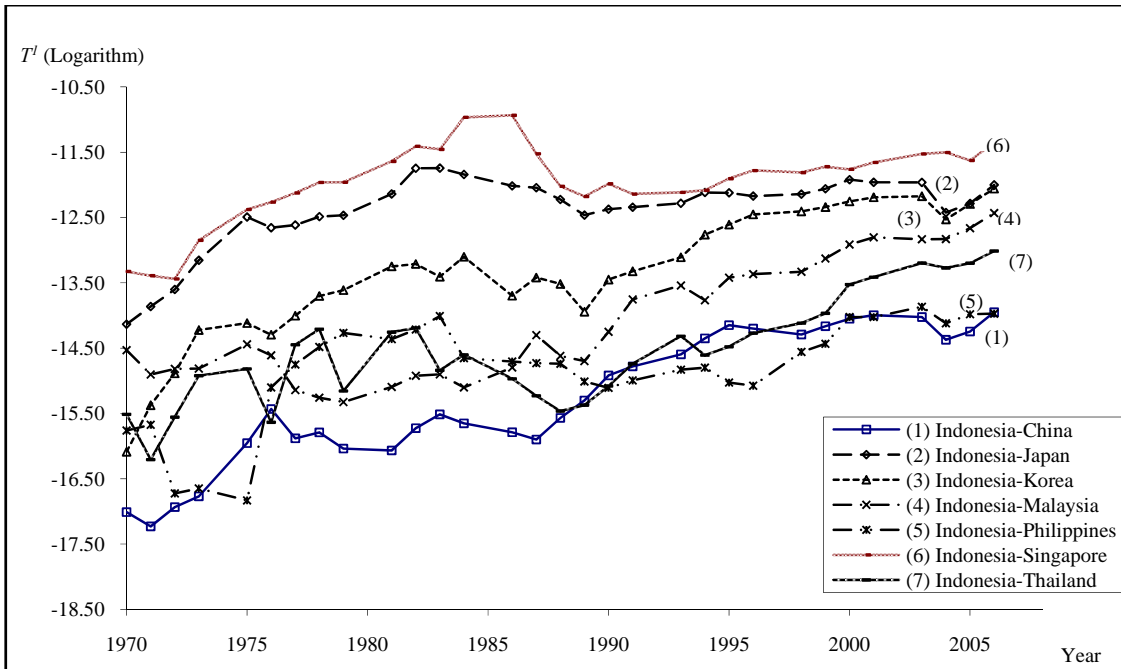


Figure B.3 Bilateral Trade Intensity  $T^i$  between Japan and East Asian Countries

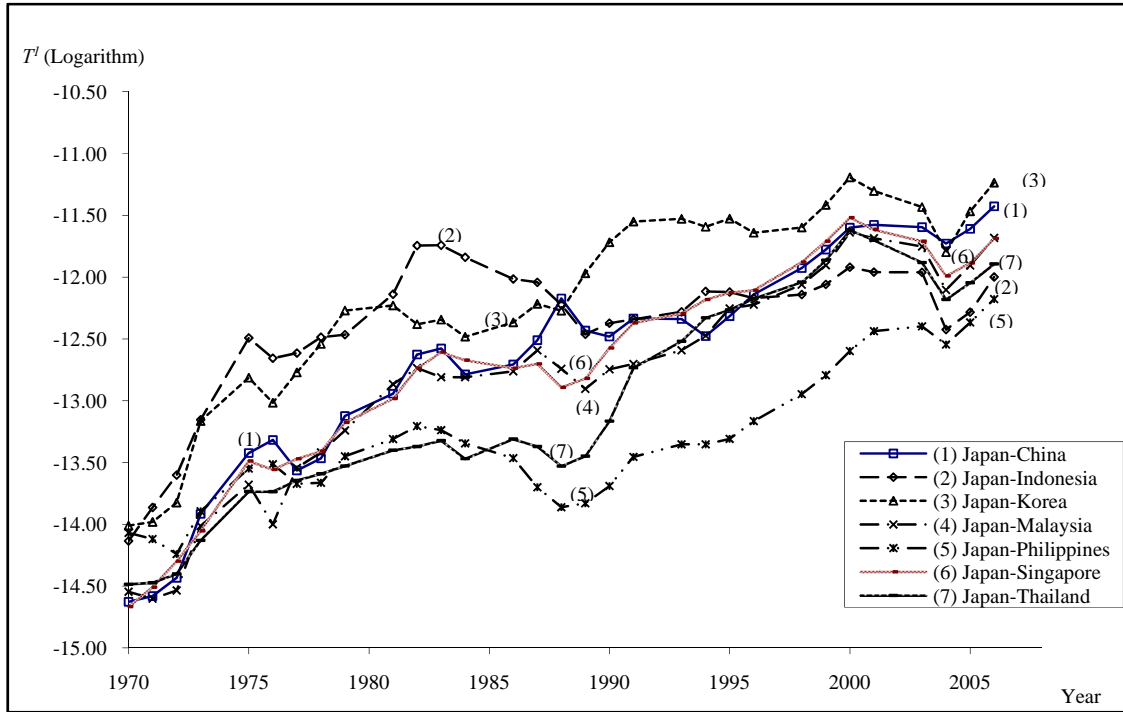


Figure B.4 Bilateral Trade Intensity  $T^i$  between Korea and East Asian Countries

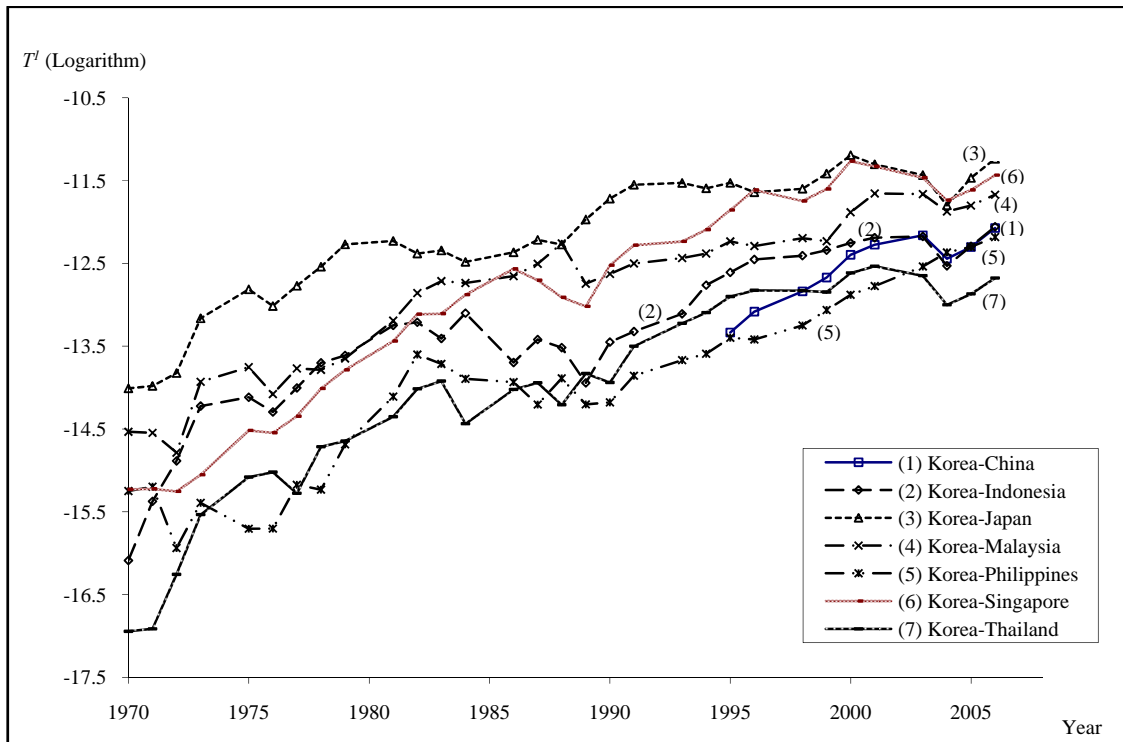


Figure B.5 Bilateral Trade Intensity  $T^d$  between Malaysia and East Asian Countries

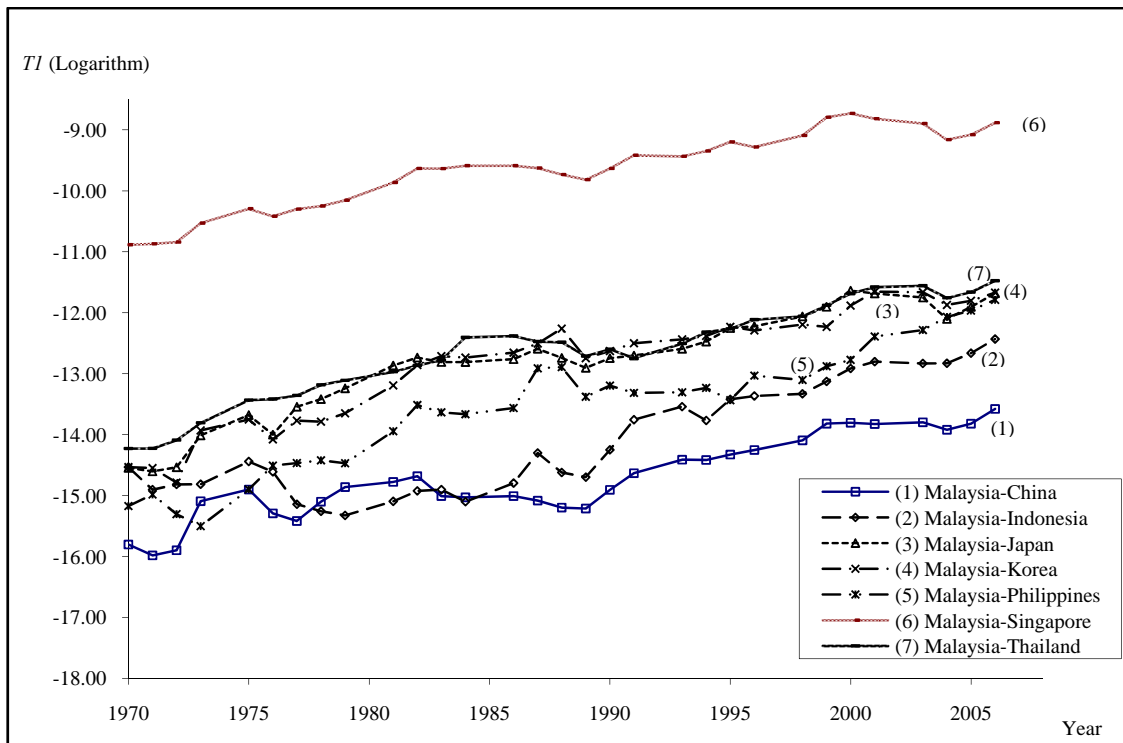


Figure B.6 Bilateral Trade Intensity  $T^l$  between the Philippines and East Asian Countries

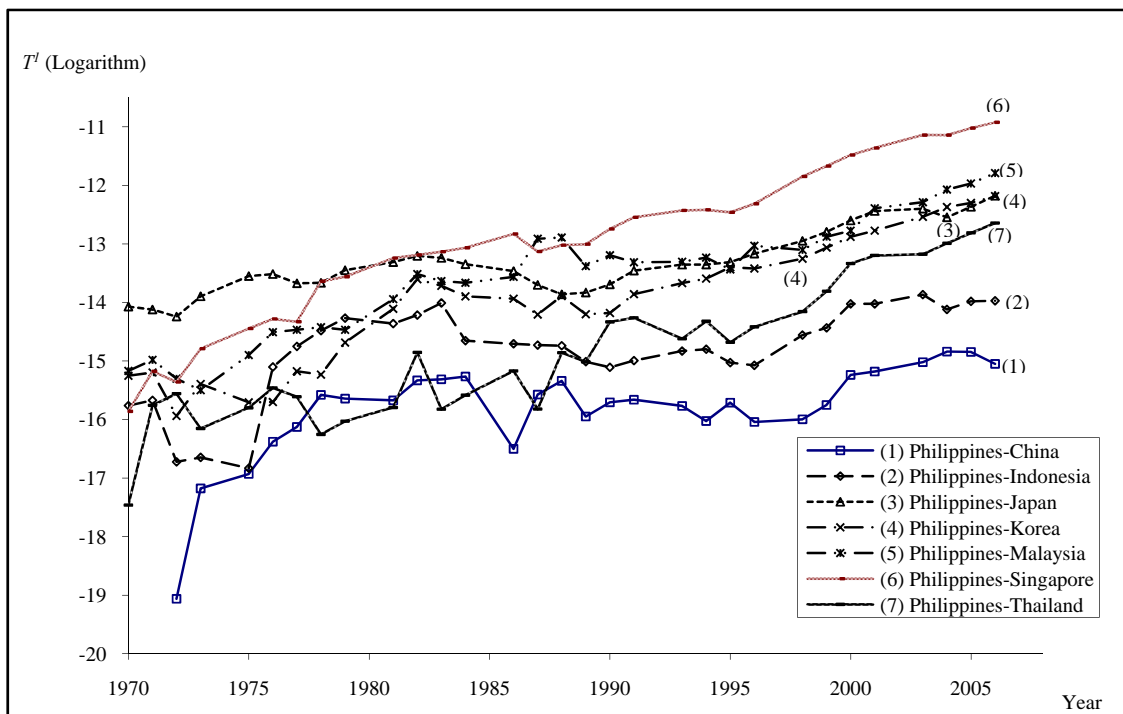


Figure B.7 Bilateral Trade Intensity  $T^l$  between Singapore and East Asian Countries

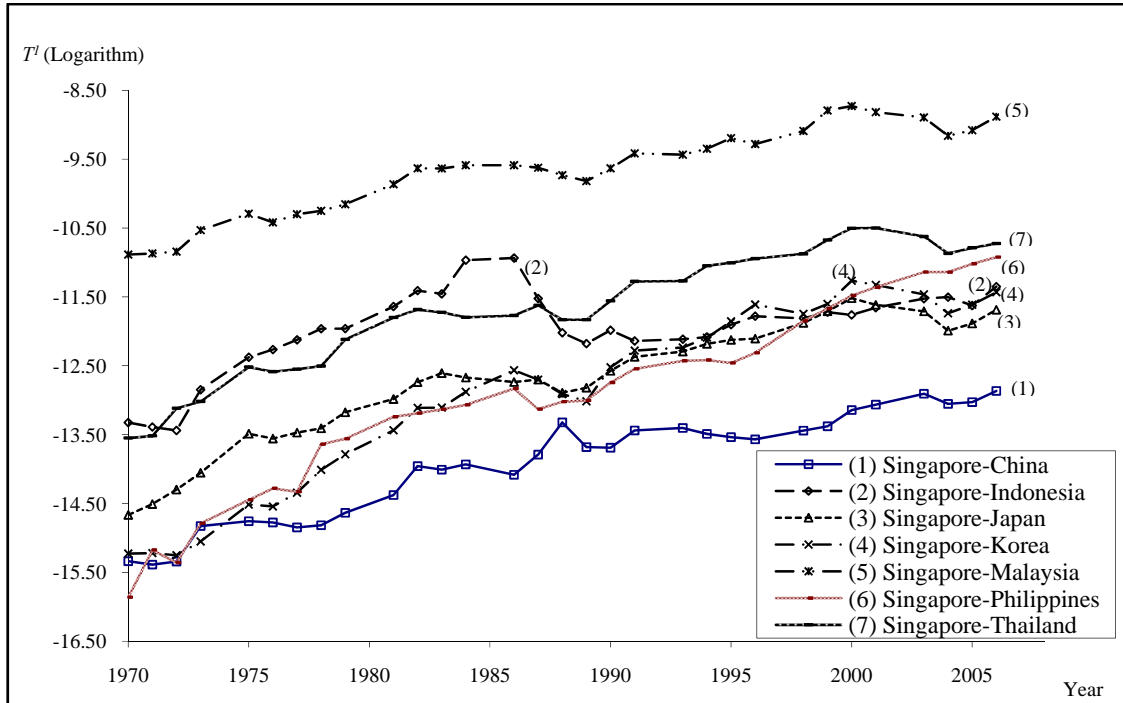
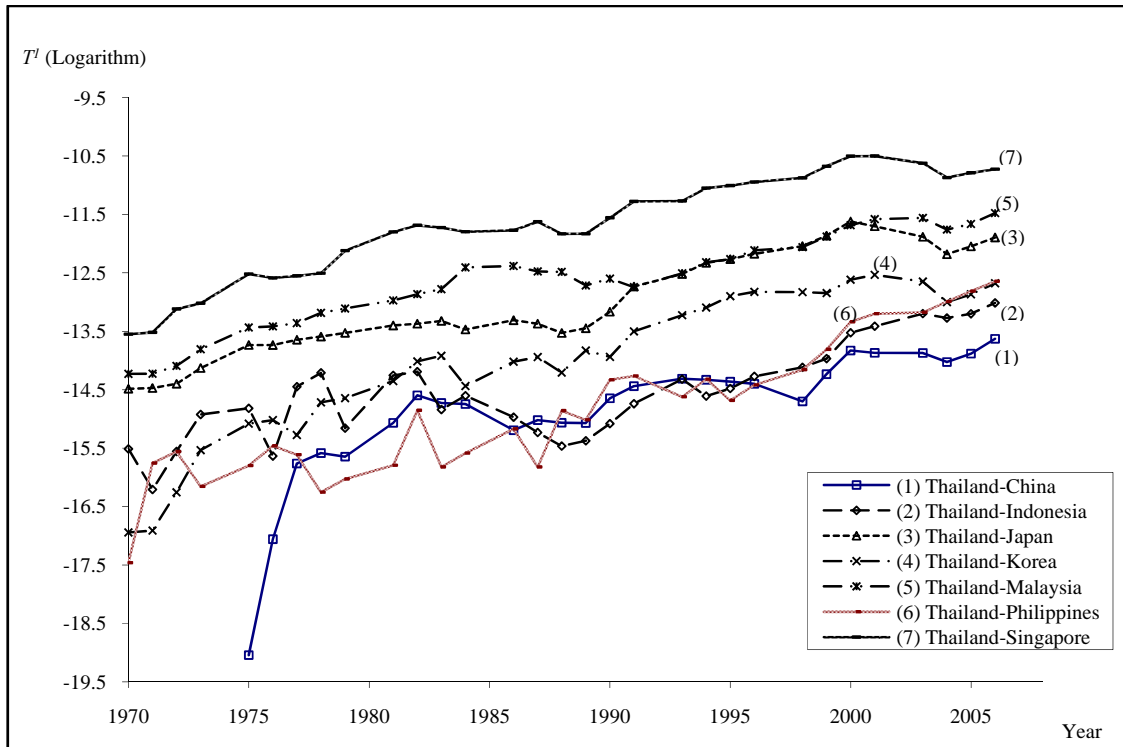


Figure B.8 Bilateral Trade Intensity  $T^l$  between Thailand and East Asian Countries



Bilateral trade intensity  $T^2$  as calculated in equation (3) for all countries are plotted here.

Figure B.9 Bilateral Trade Intensity  $T^2$  between China and East Asian Countries

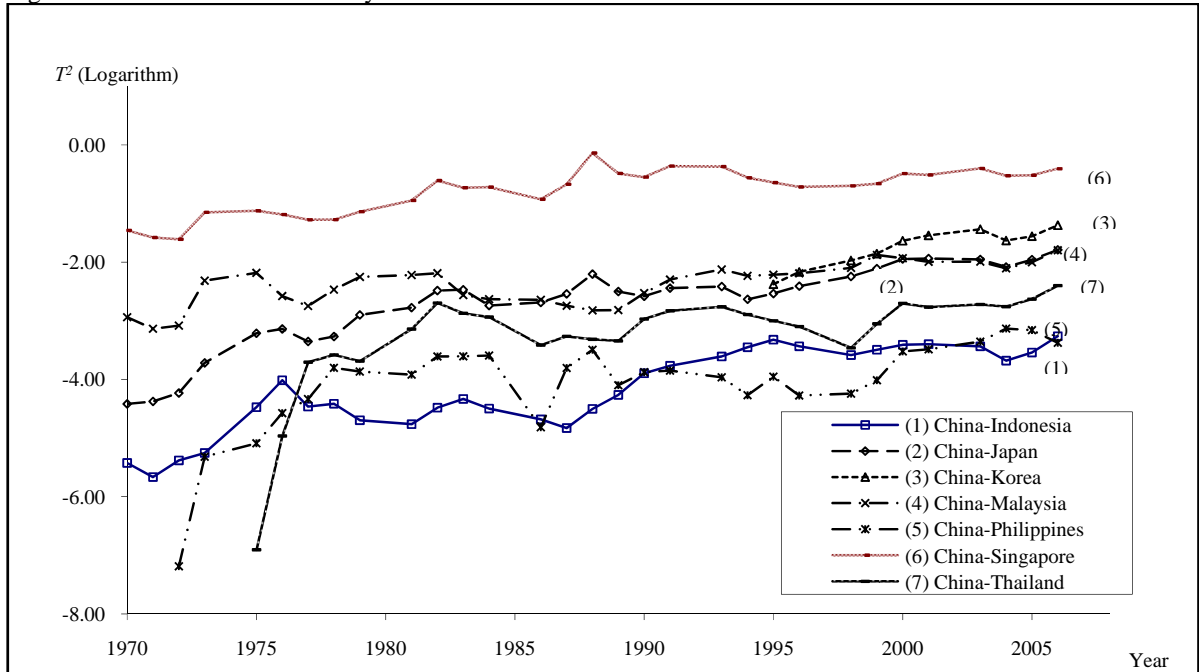


Figure B.10 Bilateral Trade Intensity  $T^2$  between Indonesia and East Asian Countries

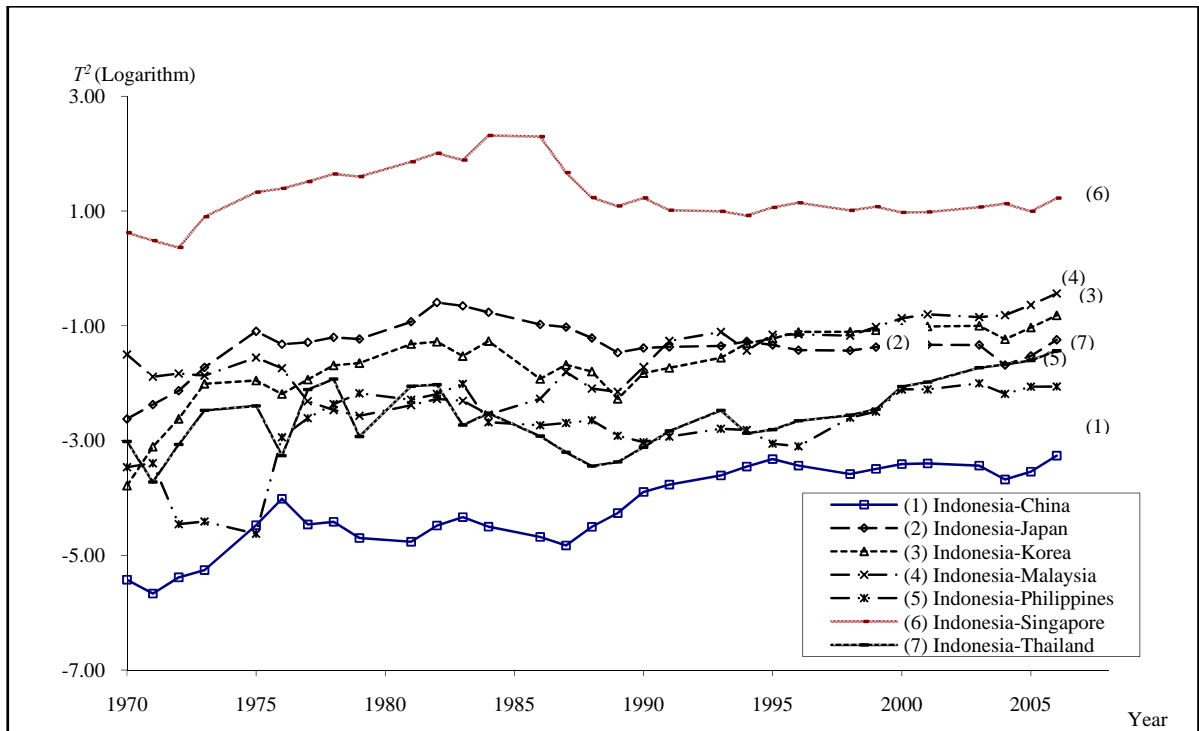


Figure B.11 Bilateral Trade Intensity  $T^2$  between Japan and East Asian Countries

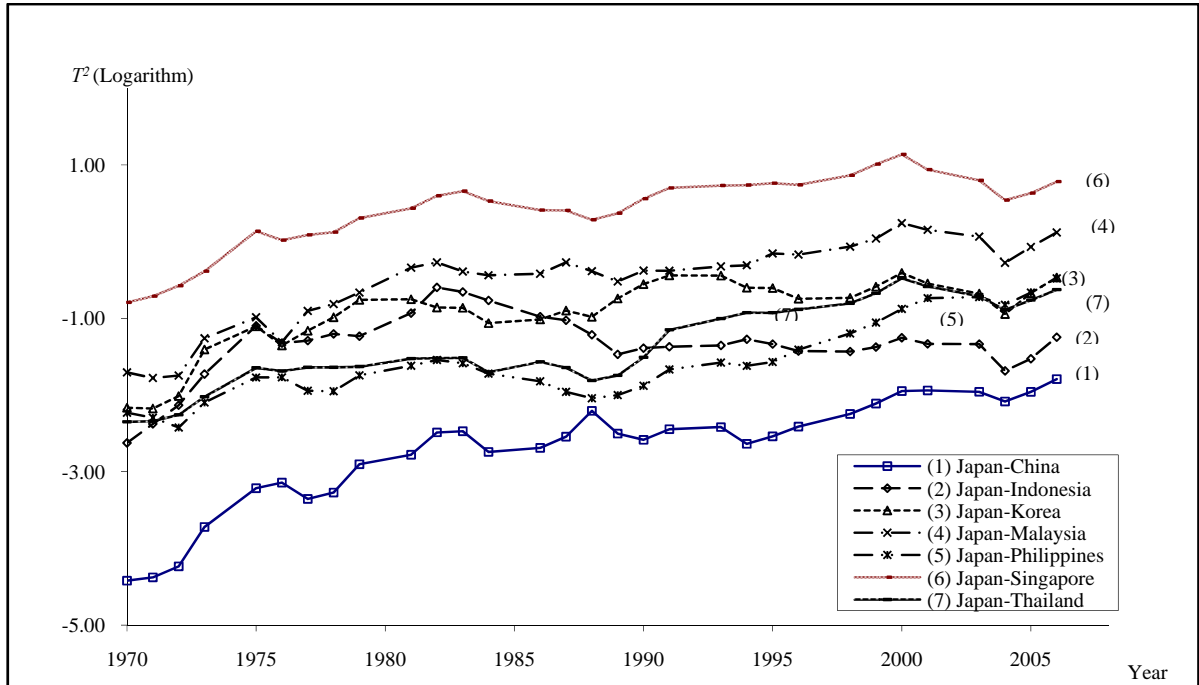


Figure B.12 Bilateral Trade Intensity  $T^2$  between Korea and East Asian Countries

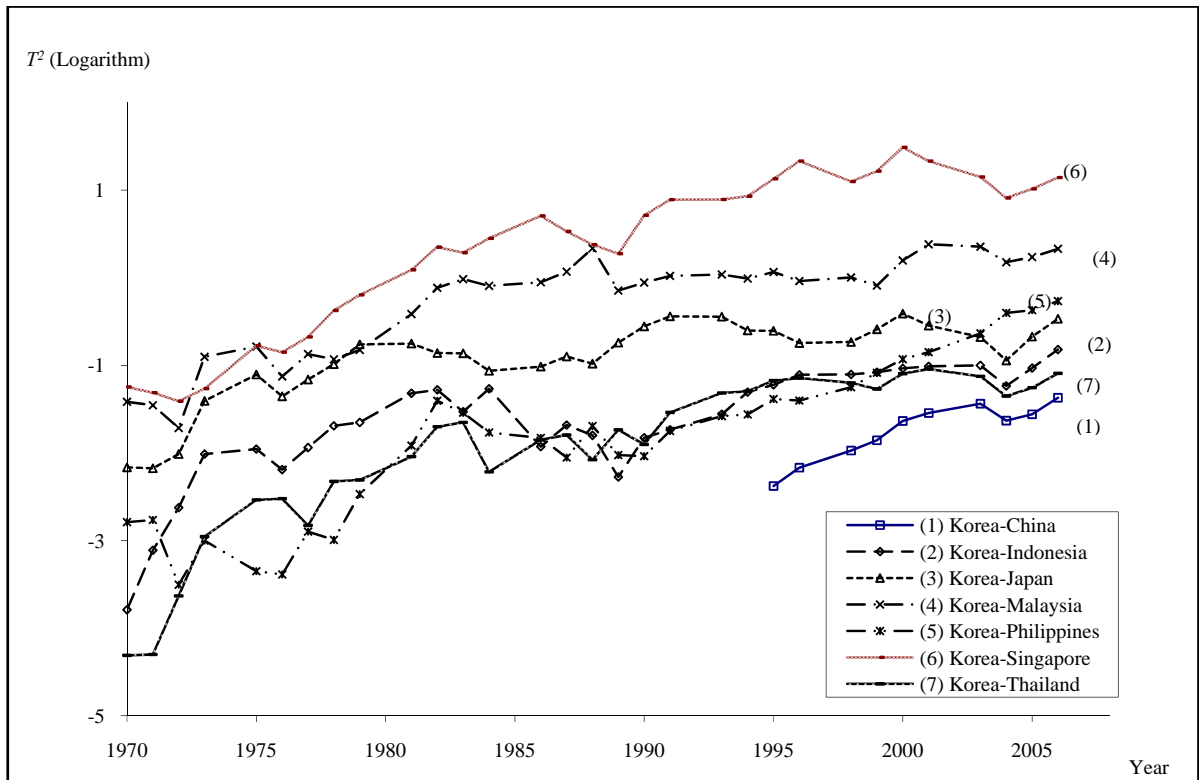




Figure B.13 Bilateral Trade Intensity  $T^2$  between Malaysia and East Asian Countries

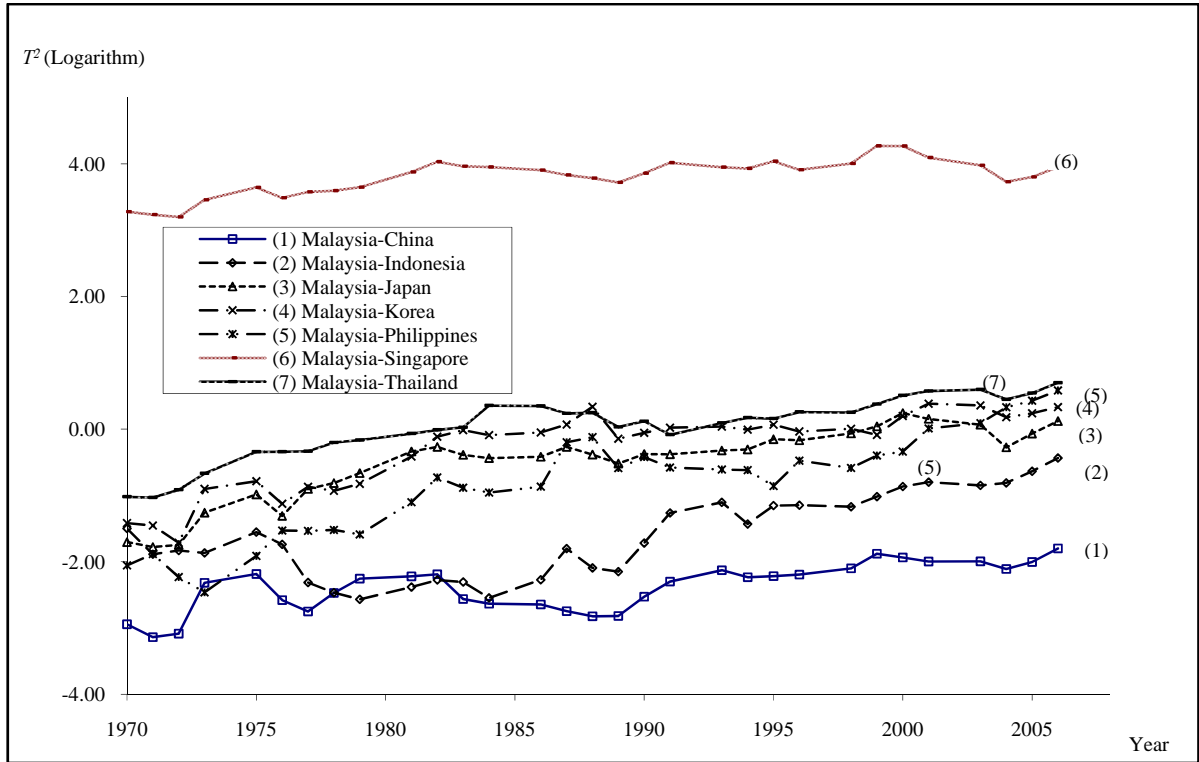


Figure B.14 Bilateral Trade Intensity  $T^1$  between the Philippines and East Asian Countries

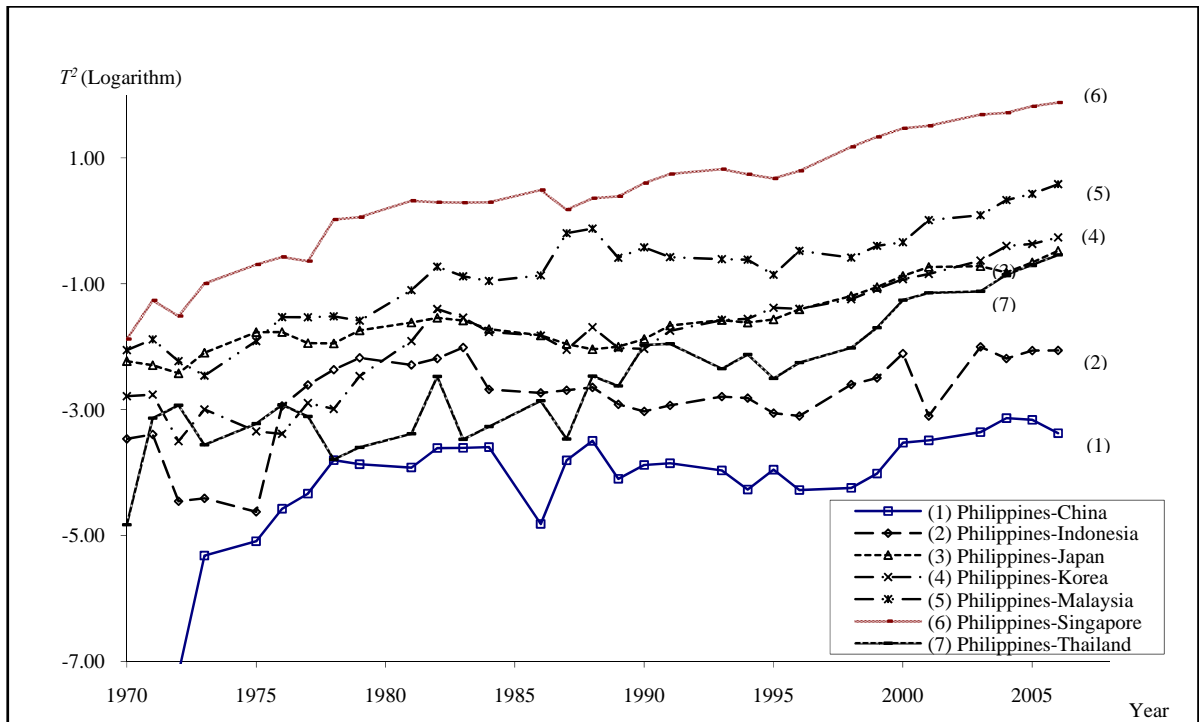


Figure B.15 Bilateral Trade Intensity  $T^l$  between Singapore and East Asian Countries

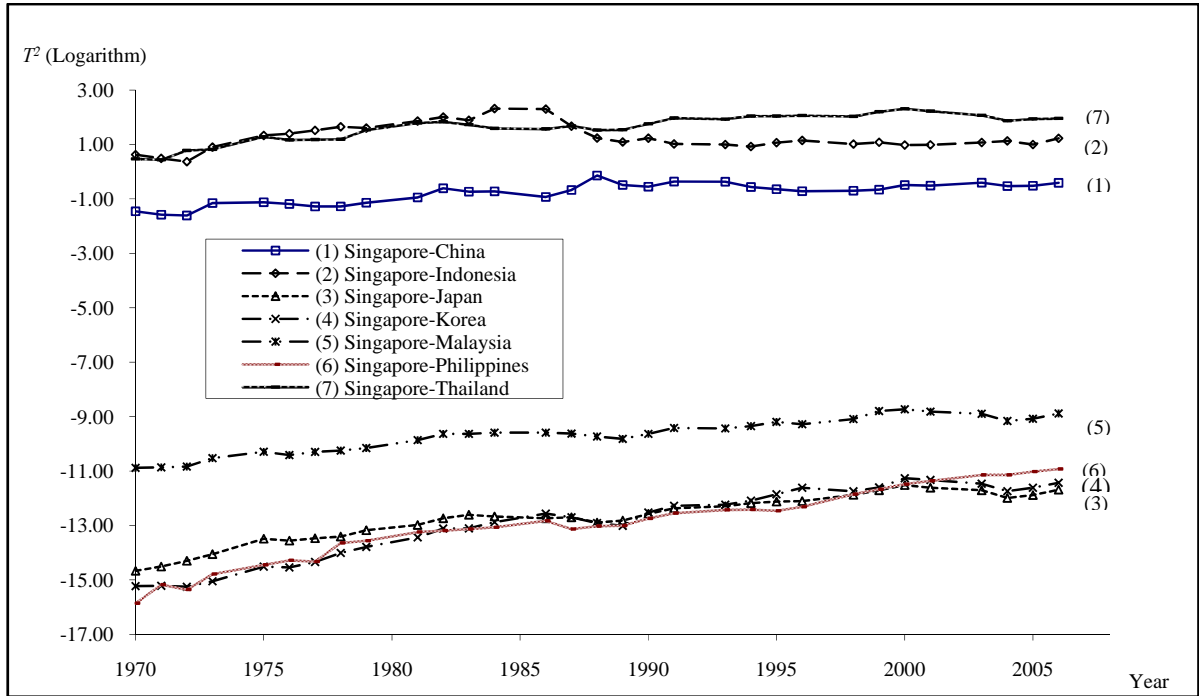
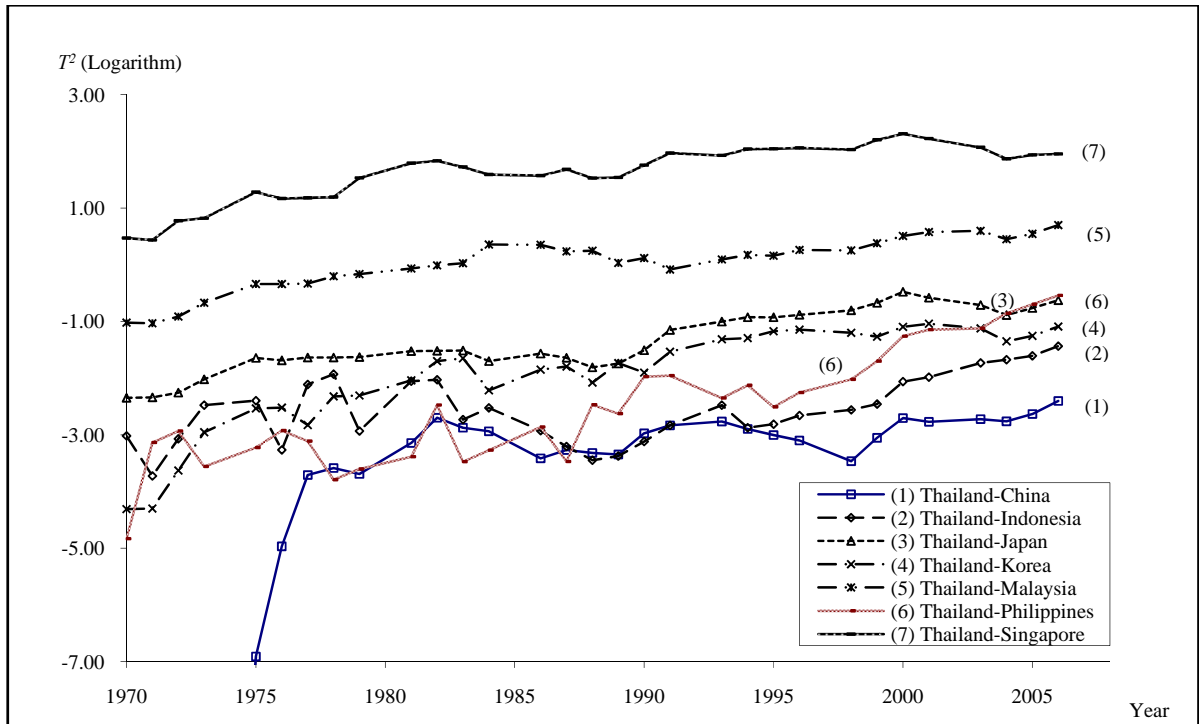


Figure B.16 Bilateral Trade Intensity  $T^l$  between Thailand and East Asian Countries



APPENDIX C

Similarity in industry specialization ( $S$ ) as calculated in equation (4) for all countries are plotted here.

Figure C.1 Similarity of Production Structures and Specialisation ( $S$ ) between China and East Asian Countries

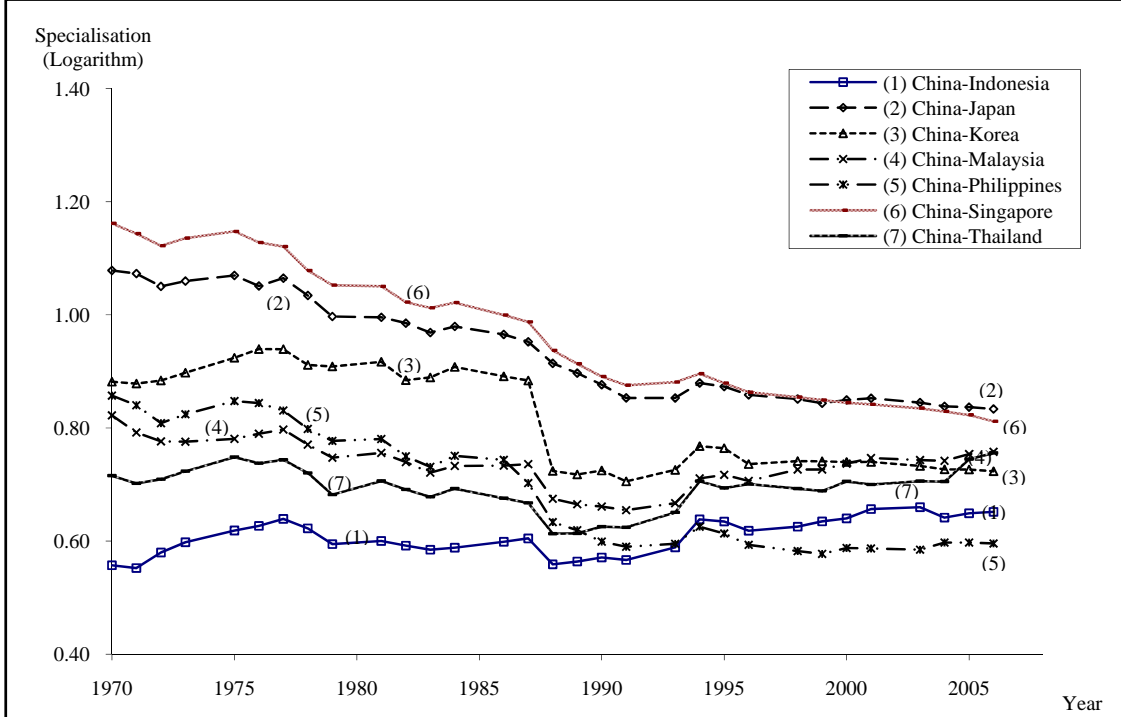


Figure C.2 Similarity of Production Structures and Specialisation ( $S$ ) between Indonesia and East Asian Countries

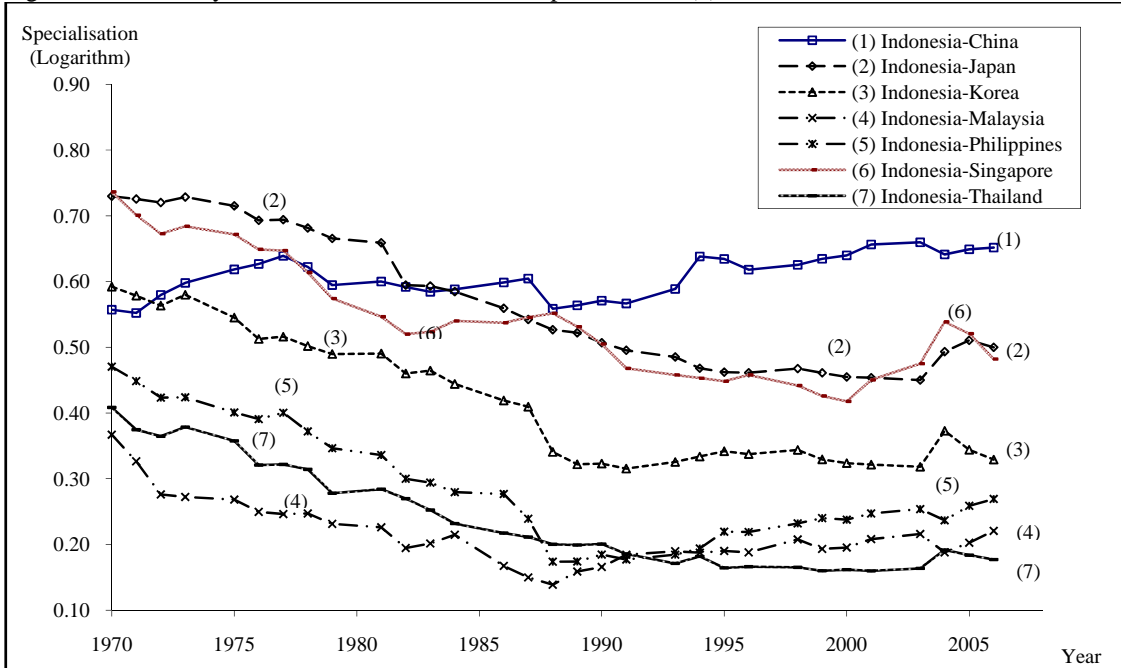


Figure C.3 Similarity of Production Structures and Specialisation (*S*) between Japan and East Asian Countries

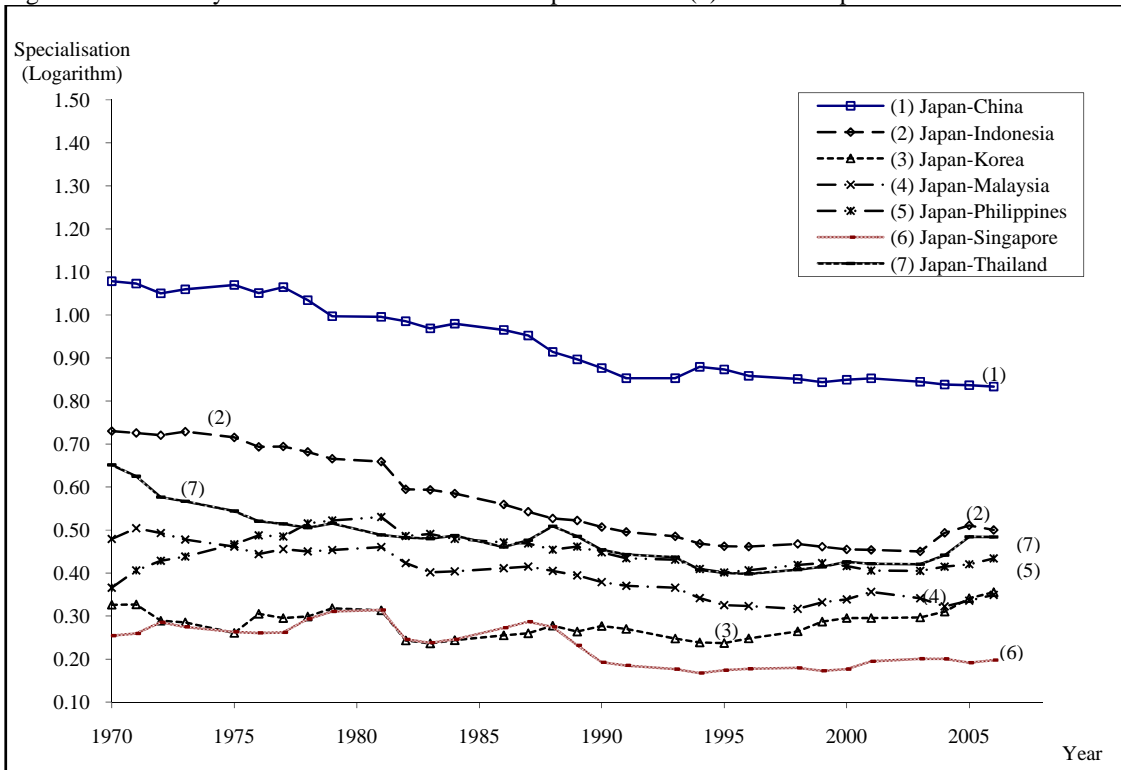


Figure C.4 Similarity of Production Structures and Specialisation (*S*) between Korea and East Asian Countries

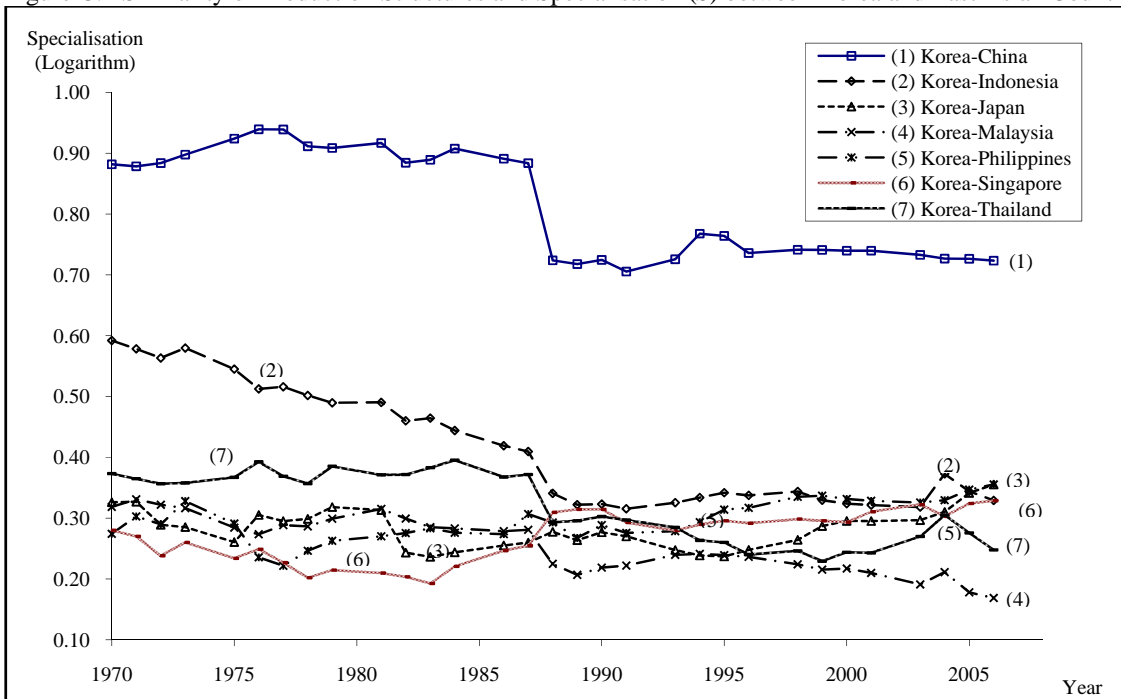


Figure C.5 Similarity of Production Structures and Specialisation (S) between Malaysia and East Asian Countries

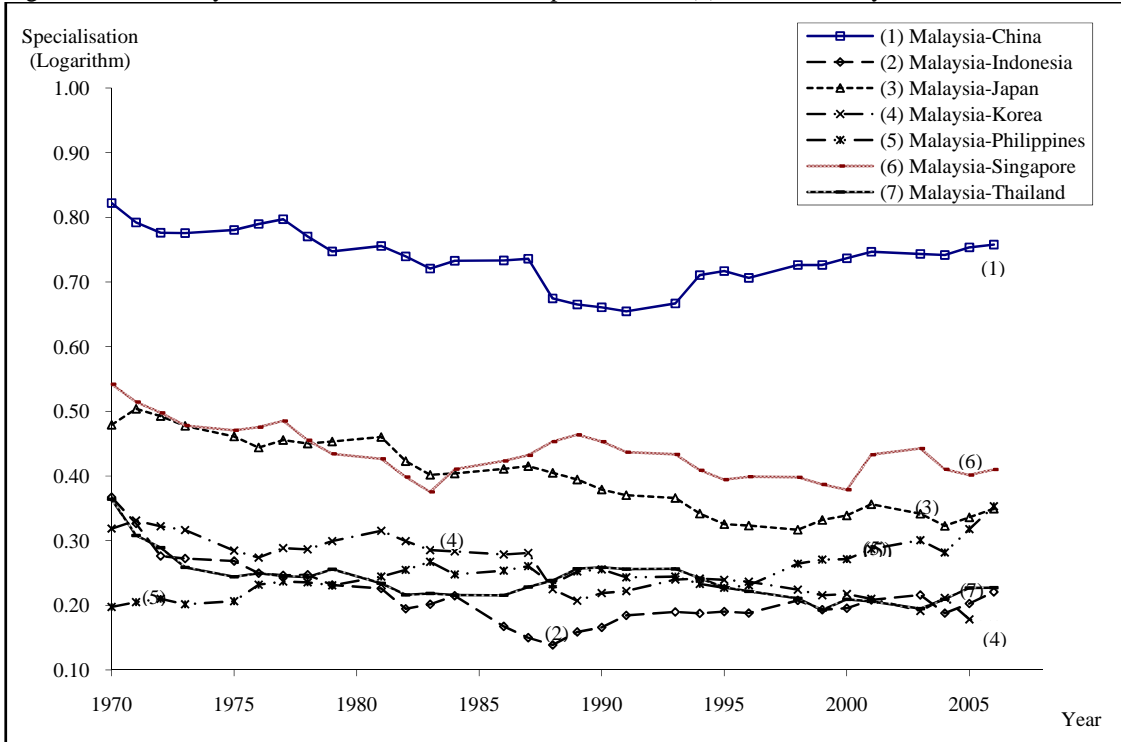


Figure C.6 Similarity of Production Structures and Specialisation (S) between the Philippines and East Asian Countries

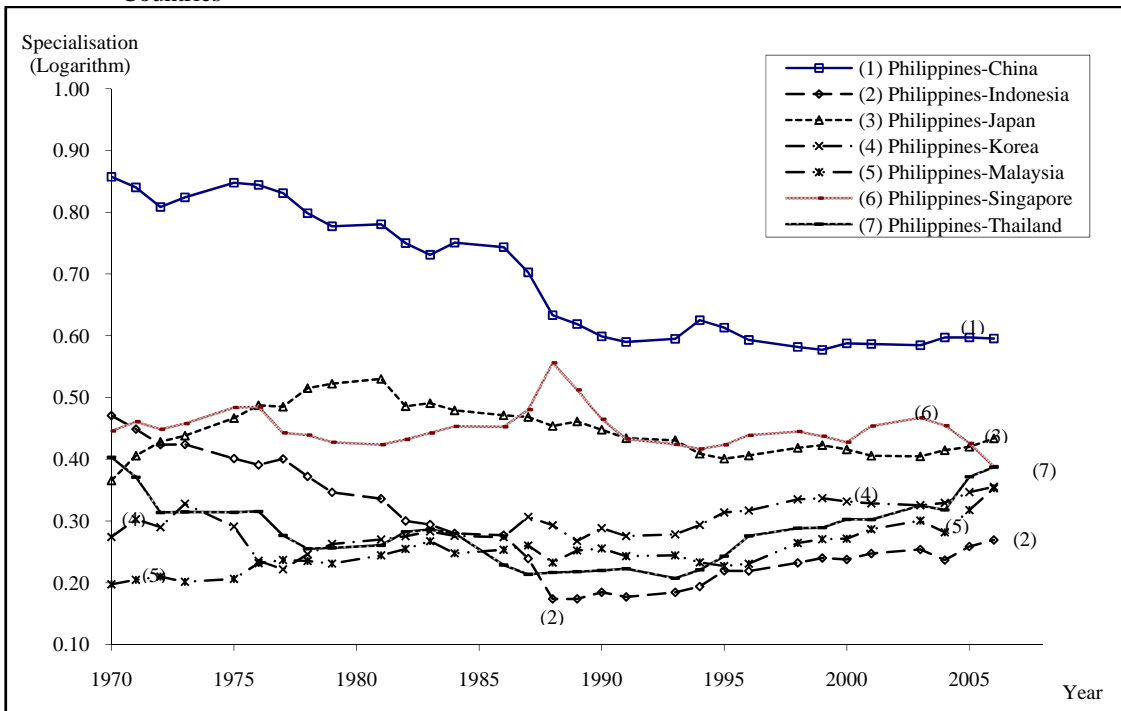


Figure C.7 Similarity of Production Structures and Specialisation (S) between Singapore and East Asian Countries

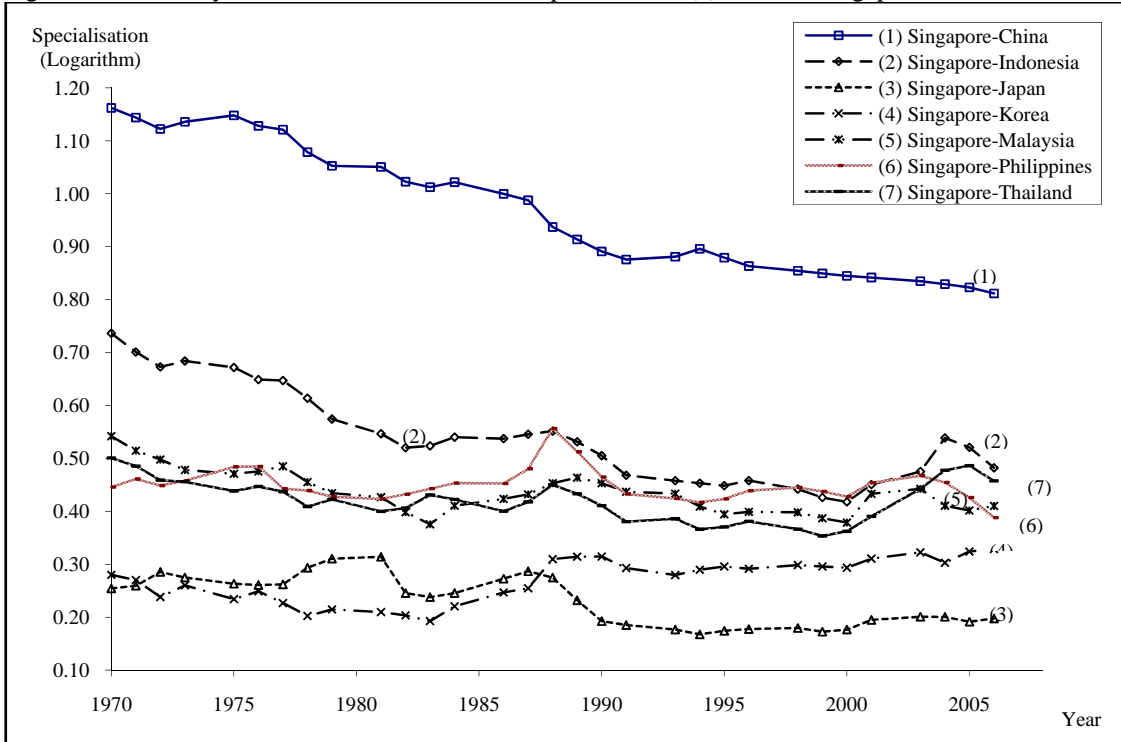
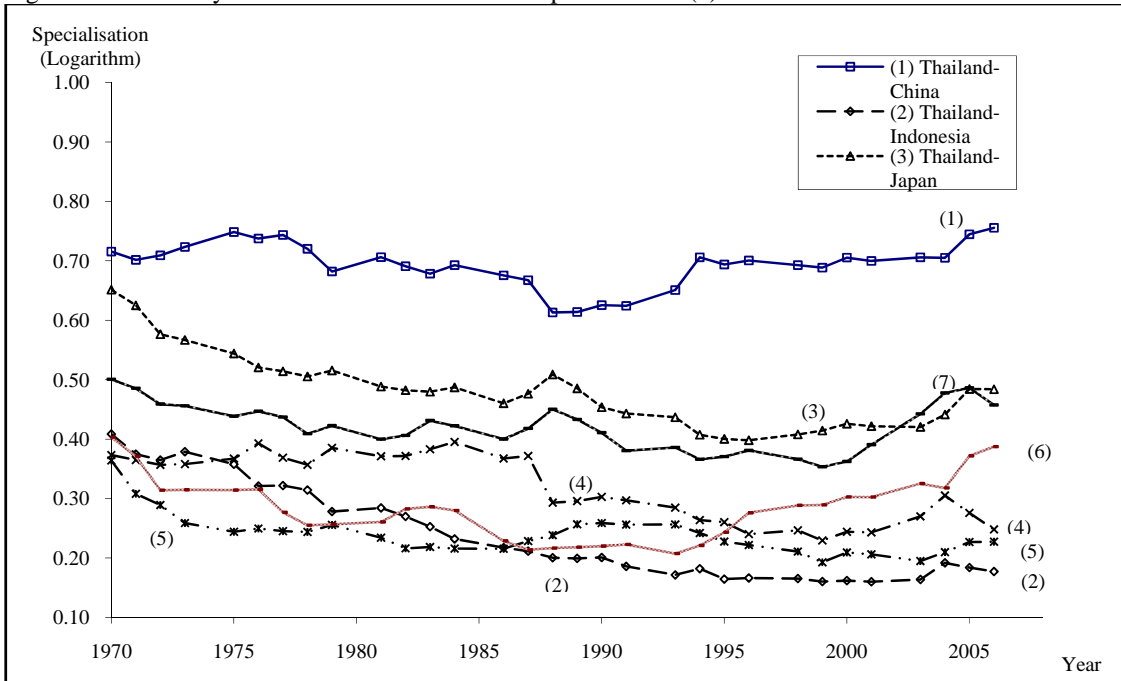


Figure C.8 Similarity of Production Structures and Specialisation (S) between Thailand and East Asian Countries



## APPENDIX D

A measure of financial integration for all twenty eight country pairs.

Figure D.1 Financial Integration Index ( $F$ ) between China and Indonesia

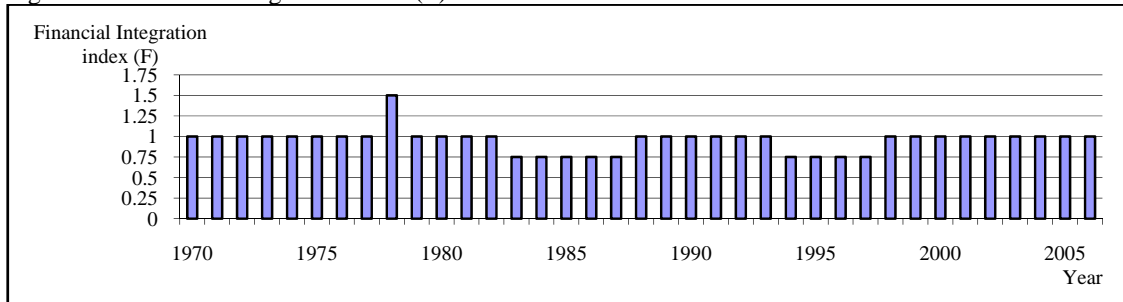


Figure D.2 Financial Integration Index ( $F$ ) between China and Japan

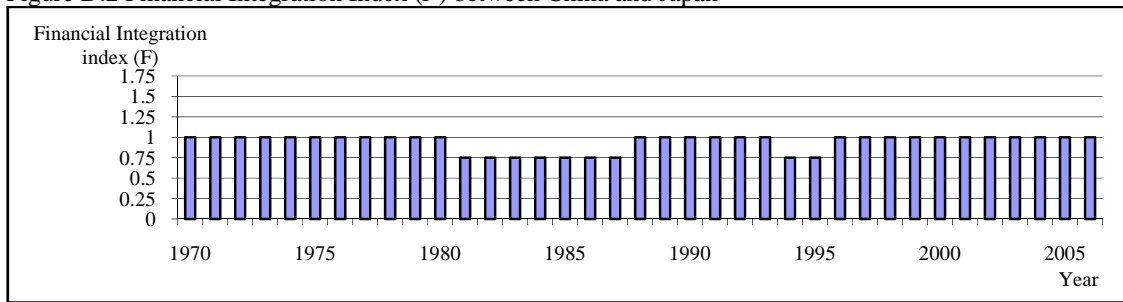


Figure D.3 Financial Integration Index ( $F$ ) between China and Korea

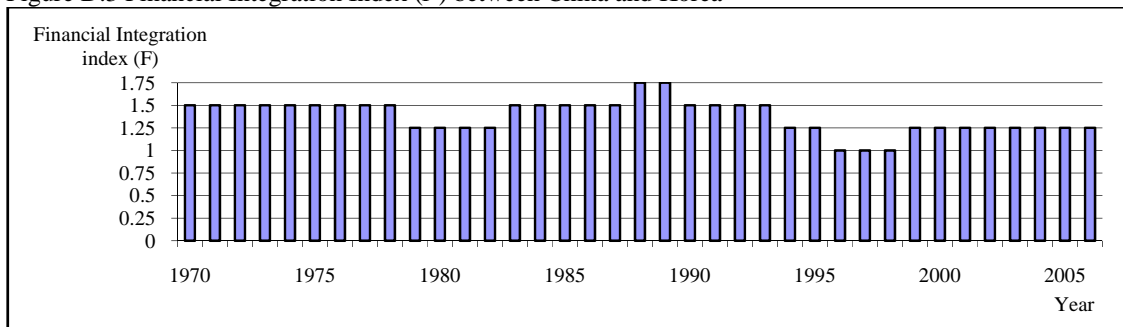


Figure D.4 Financial Integration Index ( $F$ ) between China and Malaysia

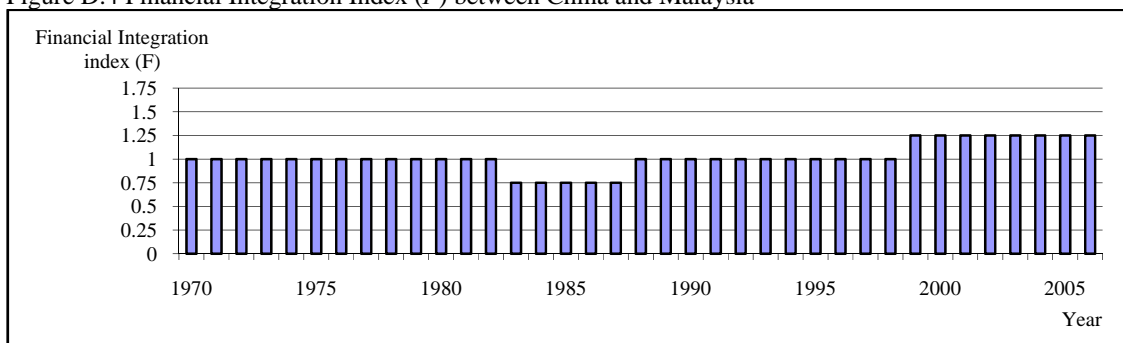


Figure D.5 Financial Integration Index ( $F$ ) between China and the Philippines

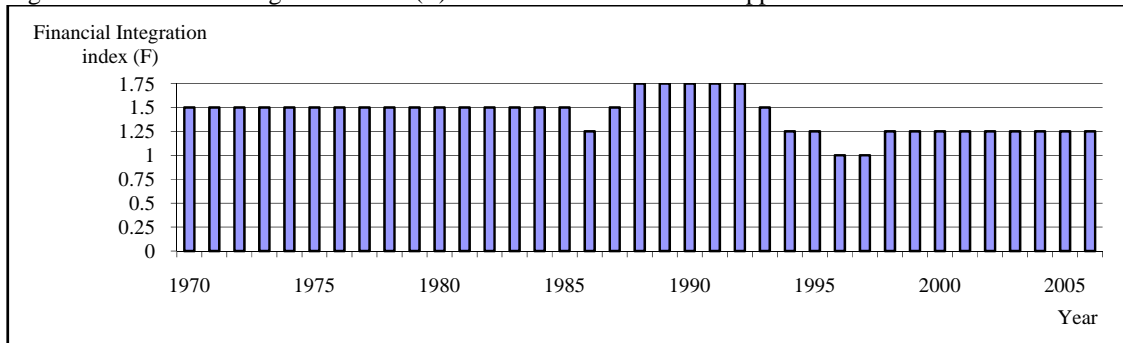


Figure D.6 Financial Integration Index ( $F$ ) between China and Singapore

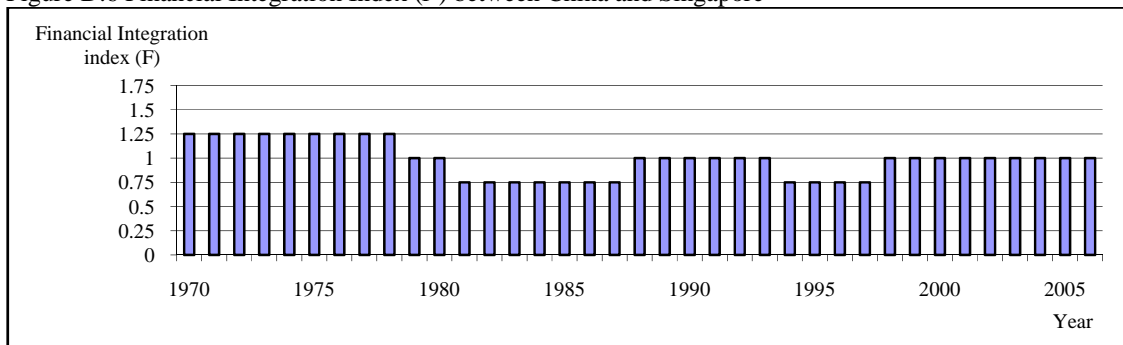


Figure D.7 Financial Integration Index ( $F$ ) between China and Thailand

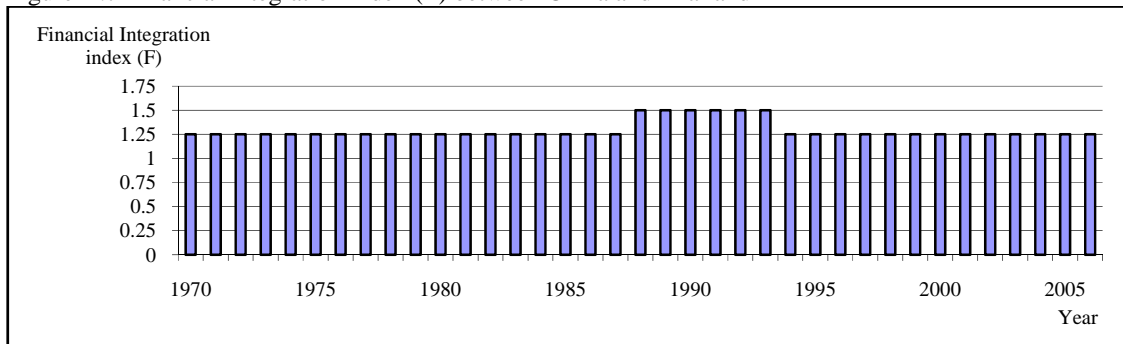


Figure D.8 Financial Integration Index ( $F$ ) between Indonesia and Japan

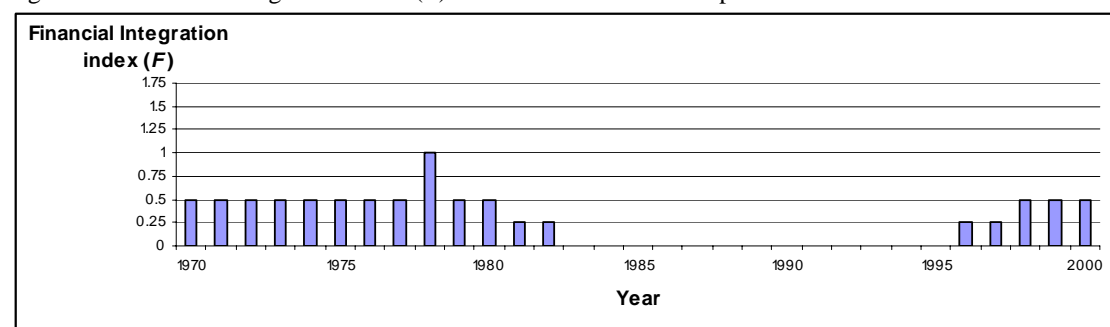




Figure D.9 Financial Integration Index ( $F$ ) between Indonesia and Korea

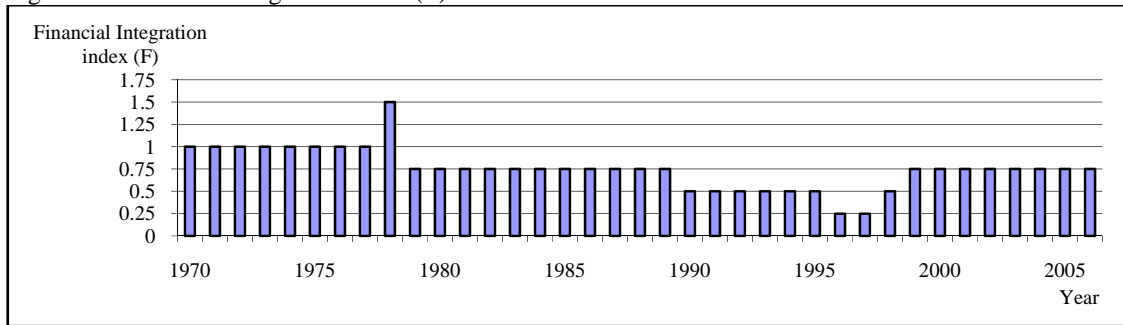


Figure D.10 Financial Integration Index ( $F$ ) between Indonesia and Malaysia

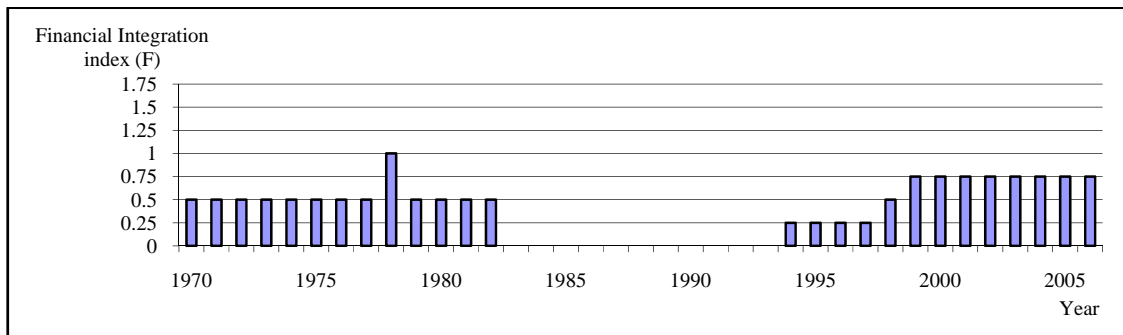


Figure D.11 Financial Integration Index ( $F$ ) between Indonesia and the Philippines

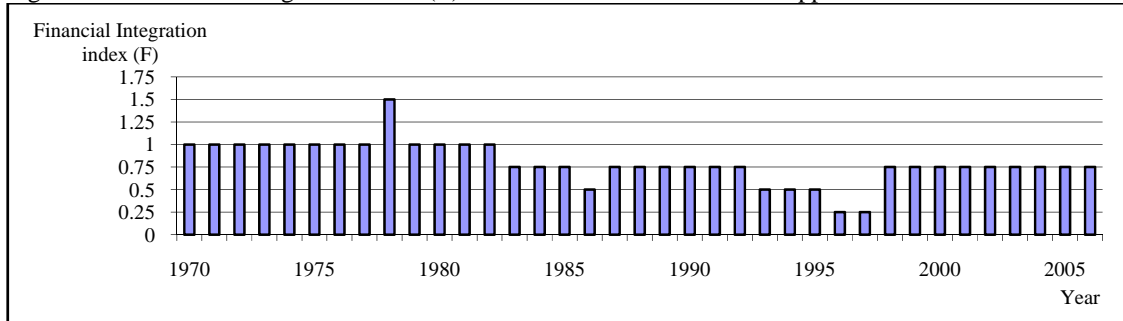


Figure D.12 Financial Integration Index ( $F$ ) between Indonesia and Singapore

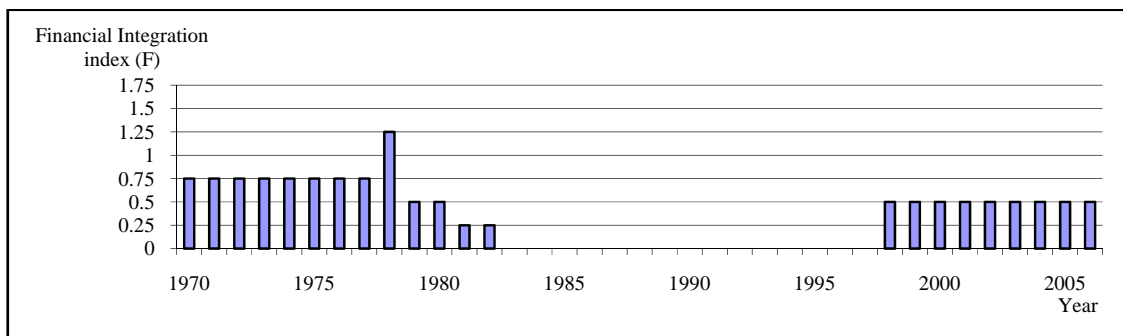


Figure D.13 Financial Integration Index ( $F$ ) between Indonesia and Thailand

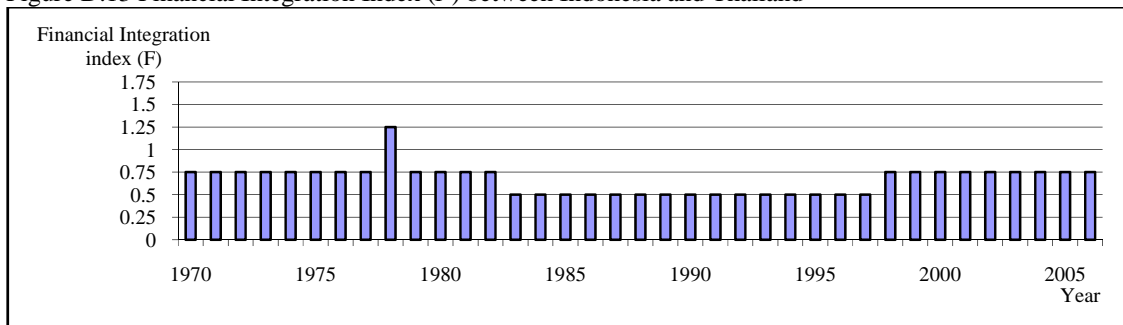


Figure D.14 Financial Integration Index ( $F$ ) between Japan and Korea

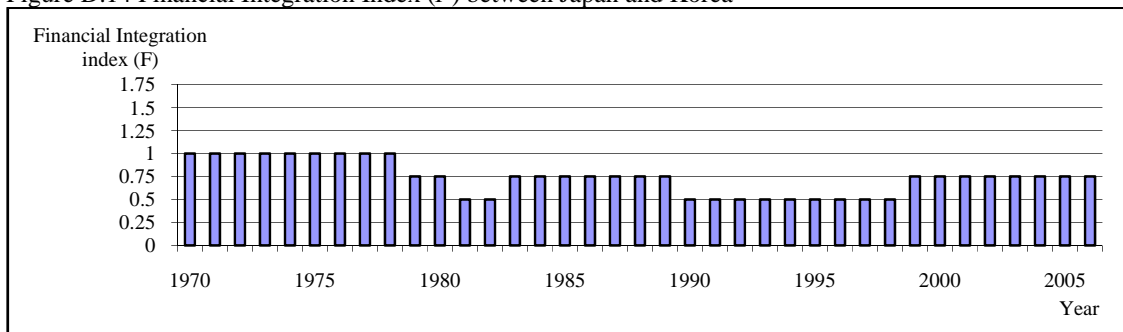


Figure D.15 Financial Integration Index ( $F$ ) between Japan and Malaysia

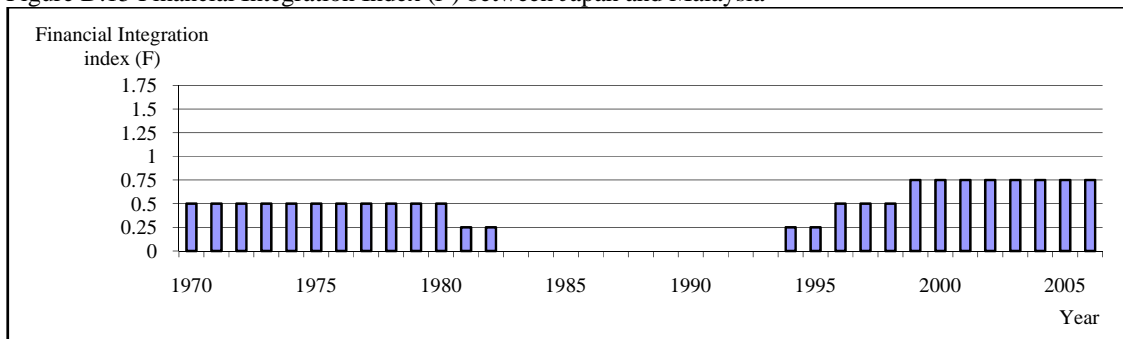


Figure D.16 Financial Integration Index ( $F$ ) between Japan and the Philippines

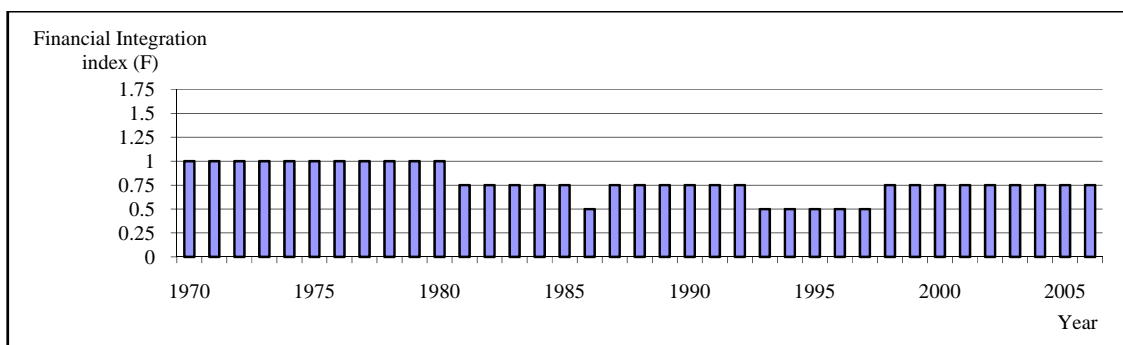


Figure D.17 Financial Integration Index ( $F$ ) between Japan and Singapore

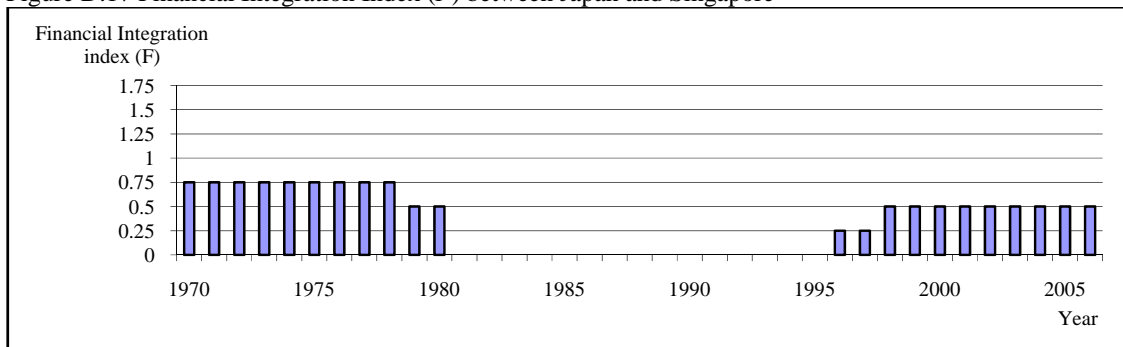


Figure D.18 Financial Integration Index ( $F$ ) between Japan and Thailand

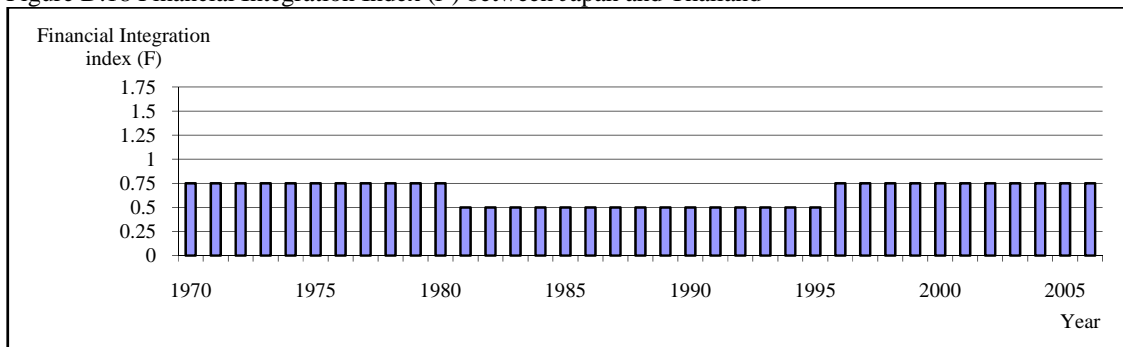


Figure D.19 Financial Integration Index ( $F$ ) between Korea and Malaysia

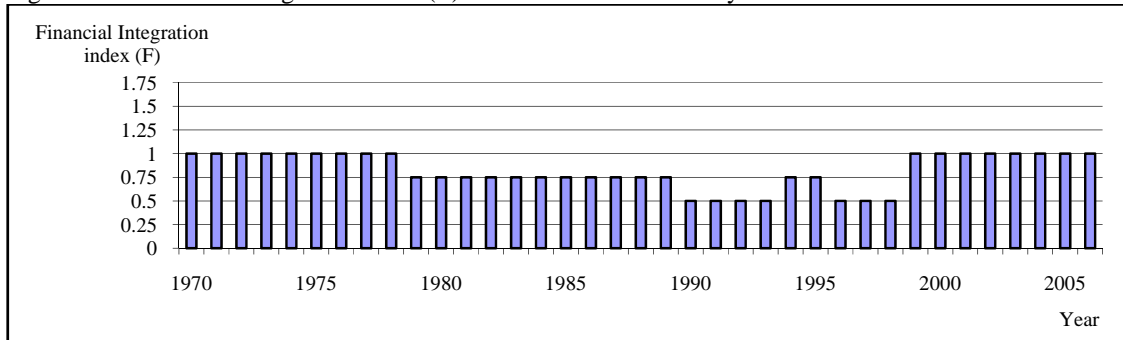


Figure D.20 Financial Integration Index ( $F$ ) between Korea and the Philippines

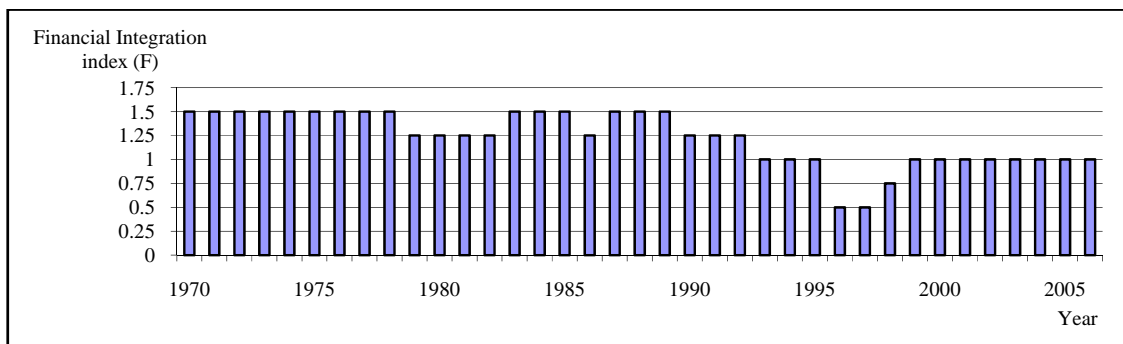


Figure D.21 Financial Integration Index ( $F$ ) between Korea and Singapore

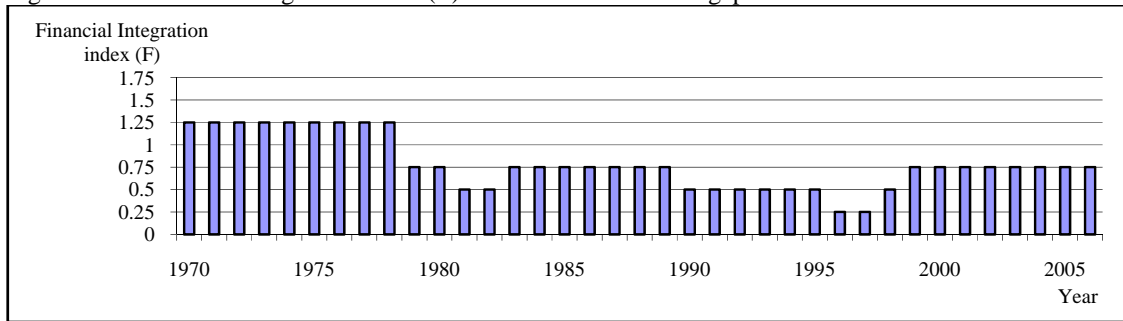


Figure D.22 Financial Integration Index ( $F$ ) between Korea and Thailand

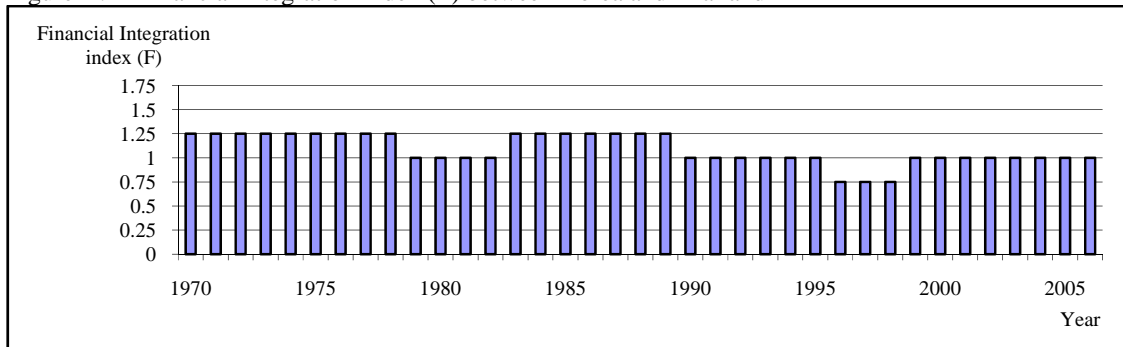


Figure D.23 Financial Integration Index ( $F$ ) between Malaysia and the Philippines

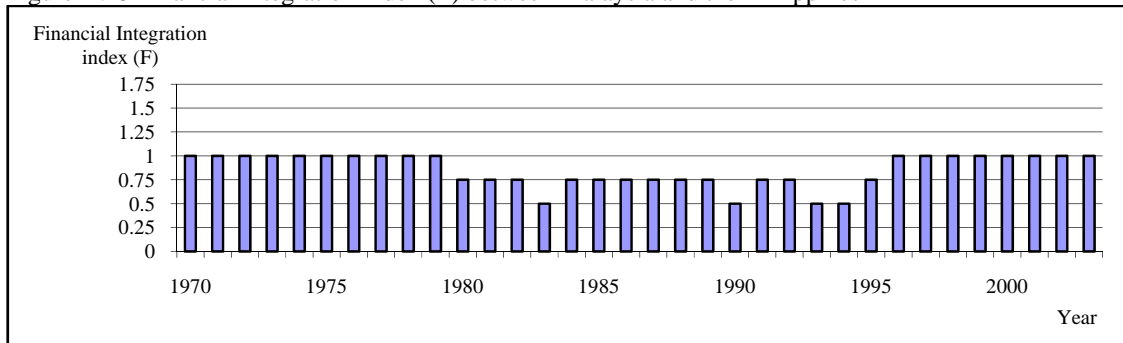


Figure D.24 Financial Integration Index ( $F$ ) between Malaysia and Singapore

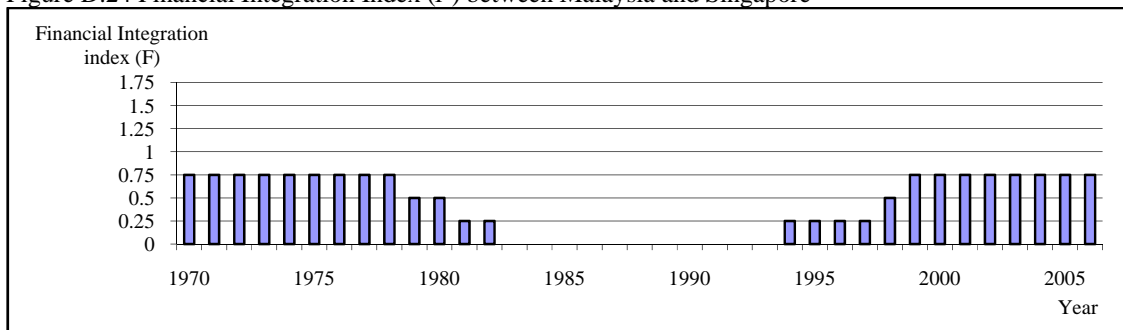


Figure D.25 Financial Integration Index ( $F$ ) between Malaysia and Thailand

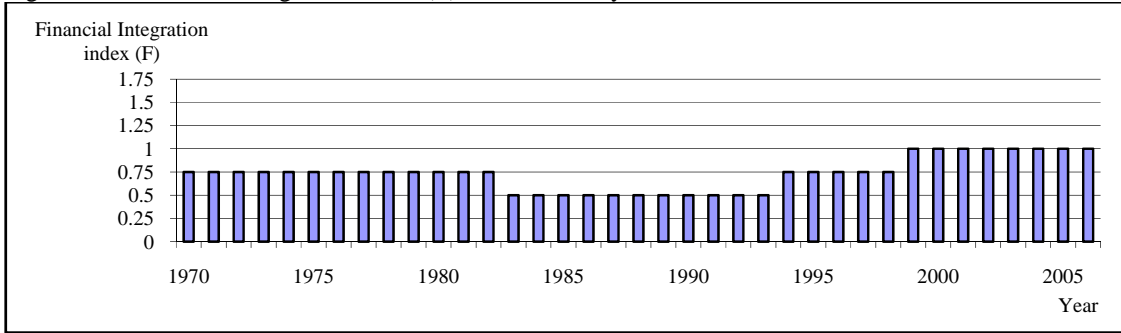


Figure D.26 Financial Integration Index ( $F$ ) between the Philippines and Singapore

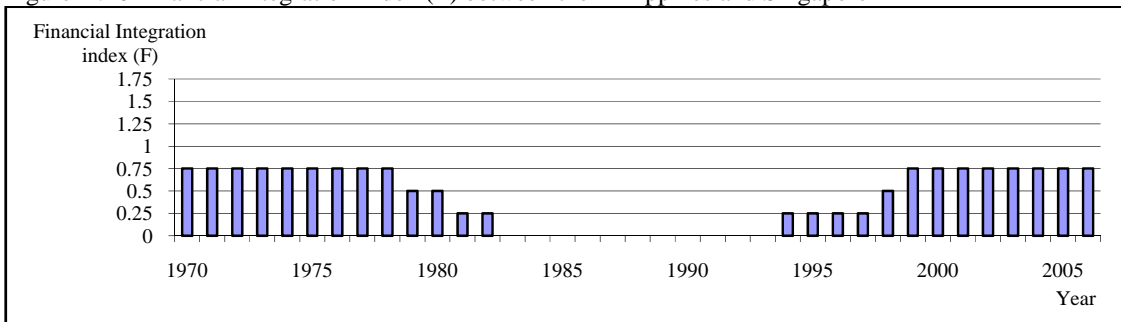


Figure D.27 Financial Integration Index ( $F$ ) between the Philippines and Thailand

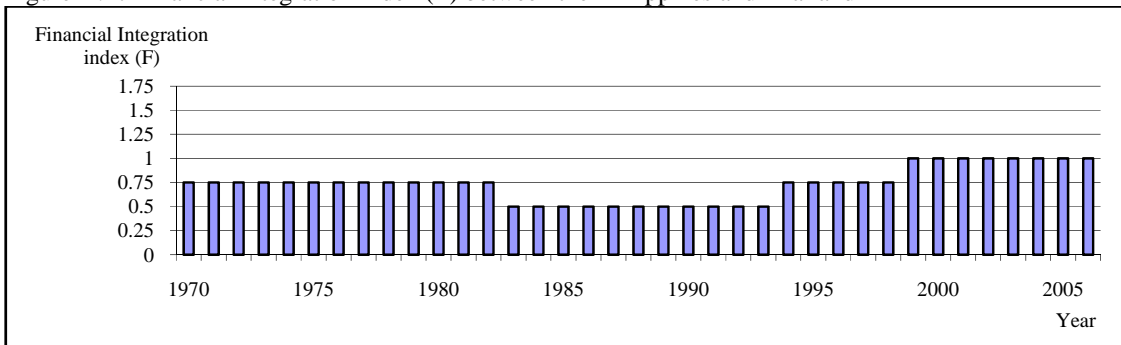


Figure D.28 Financial Integration Index ( $F$ ) between the Singapore and Thailand

