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East Asian Currency Area: A Bayesian Dynamic Factor Model Analysis

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Abstract

There has been recently increasing interest in the establishment of a common currency area in East Asia in the aftermath of the East Asian financial crisis. In this paper, we examine the desirability and feasibility of forming a currency area in the region by checking the symmetry of shocks as an important criterion of the Theory of Optimum Currency Area. We employ a Dynamic Factor Model to decompose aggregate output into global, regional and country-specific components and estimate the model using Gibbs sampling simulation. Persistent properties of those components are examined and variance decomposition analysis is performed to investigate the role of each component in output variance. Based on variance analysis, we find that East Asia countries, on average, are less plausible candidates for a currency area than European counterparts. However, a subgroup of countries in East Asia are as qualified as those in Europe. Given the ongoing integration in East Asia, it is not premature to prepare for such a currency area in this region.

Keywords: East Asia, Currency Area, Bayesian, Dynamic Factor Model, Gibbs Sampling

JEL classification numbers: F33, F42

I. Introduction

There has been a resurgence of interest in a concerted monetary arrangement and currency union in East Asia in the aftermath of the regional crisis. In both academia and policy circle, the issue of establishing a regional currency area has attracted increasing attention (see for example, Kwan (1998), Kuroda (2004)). Indeed, as an initial step, the ASEAN+3 (ASEAN plus Japan, Korea and China) has agreed, in the so-called Chiang Mai Initiative, upon a network of bilateral swap agreements which allow East Asian countries to borrow fund from each other. The issuance of Asia regional accounting currency (ACU) has been put forward.

This paper assesses the feasibility and desirability of forming a currency area in East Asia. The theory of Optimum Currency Area (OCA) has been an important guideline for assessing the possibility of a currency area. Joining a currency area typically requires abandoning the monetary independence of member countries in favor of a common policy. The desirability of a currency area depends largely on the cost of this abandonment against the benefits one can reap. The work of Mundell (1961), McKinnon (1963), Kenen(1969) and others have culminated in a host of criteria, upon which one could judge whether a country is suitable, in economic sense, for a currency union. Among these criteria, the symmetry of shocks seems to be of paramount importance. If shocks are symmetric, the need for separate monetary policies is minimal since a common policy serves all countries reasonably well. Otherwise, the cost of shock adjustment might be large and one might need an independent monetary policy as an anti-cyclical instrument, unless adjustment to shocks is fast. In this paper, we examine the symmetry of output shocks in East Asia and its persistence properties in comparison with Western Europe (hereafter referred to as Europe). If the degree of shock symmetry in East Asia appears to be close to that of Europe, it might be safe to conclude that East Asia might well constitute a currency area.

While most of studies on East Asia currency area are descriptive, few have attempted to approach the issue empirically. Bayoumi & Eichengreen (1994) apply a bivariate structural VAR (SVAR) model *à la* Blanchard & Quah (1989) to identify supply and demand disturbances underlying aggregate output and price. They compute bilateral correlations between each country disturbances with that of Japan. Higher correlation means that shocks are more symmetric and that it is more plausible to join a currency area. Their result indicates that East Asia is as suitable for a currency area as the EU members. Sato *et al* (2003) extend the SVAR model to three and five variables, incorporating real exchange rate, price and foreign output. Compared to Europe, they show that it is less suitable for East Asia as a whole to establish a currency area than has been suggested in earlier studies. Only sub-groups of East Asia countries, such as East Asia NICs and ASEAN members are possible candidates for further monetary integration as their underlying shocks are symmetric and adjustment speed to shocks is fast. Japan appears in their study as having no significant correlation in supply, exchange rate and demand shocks with other East Asian economies. Chow & Kim (2003) modify the SVAR analysis using an alternative identification scheme to break output disturbances into global, regional and country-specific shocks. They find that East Asian aggregate outputs are strongly influenced by country-specific shocks whereas the role of regional shocks is insignificant. Goto & Hamada (1994) and Goto (2002), on the other hand, implement principal component analysis to measure the degree of confluence in several macroeconomic variables in 14 East Asian countries (ASEAN 10 plus Korea, China, Taiwan, HongKong). They find that macroeconomic variables in East Asia are strongly synchronized and synchronization with Japan has been increasing since the 1990s. Accordingly, they claim that East Asia is well suited for a currency area and the matter is how to realize it. Thus, result has so far been mixed and it is not clear whether recent attempt for monetary integration is

justified.

These studies, however, have several disadvantages. First, SVAR analysis bases its judgement on bilateral correlations of disturbances and often requires a representative country as a proxy for regional shocks (Germany in Europe and Japan in East Asia, for example). Since what we are concerned with is regional rather than bilateral, judgements based on bilateral measurement might not be appropriate. Shocks in the representative country might also be poor proxies of regional ones. Moreover, in East Asia, there is no country naturally taking the role that Germany does in EU. Second, both SVAR and principal component approaches focus exclusively on correlation without distinguishing between global and regional causes. There might be the case that global shocks affect countries in a region simultaneously so as to induce a high level of correlation. What is computed as regional shocks might actually be global shocks. Joining a broader monetary area might be desirable in this case.

In this paper, we take another approach using a Dynamic Factor Model to decompose aggregate output into (unobserved) common components and idiosyncratic component. These common components and their coefficients are called factors and factor loadings respectively in the literature and we would use these terms interchangeably. Pioneered by Sargent & Sims (1977) and Geweke (1977), Dynamic Factor Model has been widely used to extract common factors from a set of economic time series, such as constructing coincident economic indicators (Stock and Watson, 1991) or world business cycle (Gregory *et al*, 1997). In our model, the aggregate outputs are broken down into world, regional and country-specific components. Our intuition is that, output volatility is the consequence of shocks induced by either global factors, regional factors or factors that are specific to the country itself. If most of volatilities are determined by regional factors, a regional common policy is sufficient to counter shocks. A common policy would,

however, fail to deal with shocks that are mainly country-specific. We conduct variance decomposition analysis to clarify the role of each component in inducing output fluctuations.

Traditionally, a dynamic factor model is cast in state-space form and estimated iteratively using standard Kalman filter and log likelihood maximization. However, the procedure appears difficult to perform when cross-session dimension grows large with a large number of unknown parameters. We do not follow this convention. Instead, we exploit the advantage of Bayesian Gibbs sampling simulation, which, though computationally heavy, allows us to work with large cross section data and large number of unknown parameters. This method is applied on a data set of 34 countries covering four regions: East Asia, Europe, North America and South America. Due to the lack of data, we limit to our study on aggregate output only and leave other aggregates for future works. Since OCA criteria are qualitative rather than quantitative, it is necessary to compare East Asian estimation with a successfully established currency area. Given the success of Euro zone formation, Europe is a natural benchmark for our assessment.

This approach has several advantages over the previous researches. First, since our method does not base on bilateral correlations but on the composition of output variance, our measure better captures common shocks shared by countries in a region. No representative country is needed. Second, the dynamic factor framework allows us to introduce the global factor into the model to account for global shocks. By separating global and regional factors, we are able to take care of pure regional shocks which would have been mixed up with global shocks if we applied SVAR procedure. Lastly, in our framework, analysis of the dynamics of global and regional business cycles is possible. In this paper, we collate the cycles with historical facts and investigate the their persistence properties.

To anticipate the result, we find that East Asia as a whole is a less plausible candidate for a

currency area than Europe. Specifically, shocks in East Asia are less symmetric than in Europe indicated by less share of output variance explained by regional component. Also, for country-specific shocks, East Asia seems to adjust more slowly, suggesting higher cost for the abandonment of monetary autonomy. However, a subgroup of East Asia countries appear more synchronized than the European average, implying that these countries are suitable for a currency union and could proceed to form one in the first place. Interestingly, Japan and China, despite their economic power, are 'outliers' with little synchronization with other countries in East Asia.

The rest of the paper is organized as follow. Section II describes the model, estimation technique and data processing. Section III analyzes estimation result and performs variance decomposition analysis. The final section, as usual, is for conclusion.

II. The Model

Our model is built on several assumptions. First, we assume that aggregate output could be decomposed into world component, regional component and country-specific component. World component represents global shocks that affect virtually all countries. Typical examples of global shocks are two oil price shocks in mid 1970s and early 1980s, causing global economic recessions. Regional component, similarly, embodies regional shocks affecting simultaneously on countries inside the region but are relatively innocuous to outside countries. The East Asian financial crisis and the unification of Germany are examples of regional shocks in East Asia and Europe respectively. Country-specific component represents shocks that occur and influence within a country due to its own structural or institutional causes. Further, we assume that these components are contemporaneously uncorrelated. This assumption is necessary for our model to be identifiable. The model covers four main regions: East Asia, Europe, North America and South America. Global shocks would be shared by all four regions while regional shocks are specific to a region

only. Both global shocks and regional shocks influence differently in different countries, as indicated by corresponding coefficients. The model is as follow:

$$y_{i,t} = \alpha_i W_t + \beta_i R_t^r + \varepsilon_{i,t}, \quad (t=1,2,\dots,T, \quad r=1,2,3,4, \quad i=1,2,\dots,n) \quad (1)$$

where $y_{i,t}$ stands for aggregate output of country i ; W_t stands for the world component and $R_{r,t}$ is region r component ($r = 1$ for East Asia, $r = 2$ for Europe, $r = 3$ for North America and $r = 4$ for South America). n is the number of countries and T is sample size. The error term $\varepsilon_{i,t}$ is considered as country-specific component in country i . W_t , R_t^r and $\varepsilon_{i,t}$ are contemporaneously uncorrelated. α_i and β_i are coefficients (or factor loadings of the factors W and R) measuring relative impact of the world and regional components on country i aggregate output.

Second, we assume that the components follow stationary univariate first-order autoregressive representation:

$$W_t = a W_{t-1} + \eta_t^w, \quad \eta_t^w \sim N(0, \sigma_w^2) \quad (2)$$

$$R_t^r = b_r R_{t-1}^r + \eta_t^r, \quad \eta_t^r \sim N(0, \sigma_{R^r}^2) \quad (3)$$

$$\varepsilon_{i,t} = c_i \varepsilon_{i,t-1} + \eta_{i,t}, \quad \eta_{i,t} \sim N(0, \sigma_i^2) \quad (4)$$

Where η_t^w , η_t^r and $\eta_{i,t}$ are serially and contemporaneously uncorrelated Gaussian disturbances with corresponding variances: σ_w^2 , $\sigma_{R^r}^2$ ($r = 1, 2, 3, 4$), σ_i^2 ($i = 1, 2, \dots, n - 1, n$).

Since this paper focuses on the cyclical pattern of the components, it is reasonable to assume them to be stationary (as we show later in data processing section, we first difference aggregate output and test for stationarity using the popular ADF test). For simplicity and for saving the degree of freedom, I assume the components are first-order autoregressive. It should be noted that the model, in principle, works well for the general case of $AR(p)$ process.

With the above assumptions, it is straightforward to cast the system (1)-(4) into state space form:

$$\xi_t = F\xi_{t-1} + \nu_t \quad (5)$$

$$y_t = H\xi_t \quad (6)$$

where ξ_t is a $(n + 5) \times 1$ state vector and $\xi_t = (W_t, R_t^1, R_t^2, R_t^3, R_t^4, \varepsilon_{1,t}, \varepsilon_{2,t}, \dots, \varepsilon_{n,t})'$, y_t is a $n \times 1$ observation vector and $y_t = (y_{1,t}, y_{2,t}, \dots, y_{n,t})'$; F and H are relevant coefficient matrices; ν_t is a $(n + 5) \times 1$ state disturbance vector. $\nu_t = (\eta_t^w, \eta_t^1, \eta_t^2, \eta_t^3, \eta_t^4, \eta_{1,t}, \eta_{2,t}, \dots, \eta_{n,t})'$ and

$$\nu_t \sim N(0, Q) \quad (7)$$

where Q is a diagonal variance-covariance matrix.

For the ease of reference, we denote the stacked state vectors: $\tilde{\xi} = (\xi_1, \xi_2, \dots, \xi_T)'$, $\tilde{Y} = (y_1, y_2, \dots, y_T)'$. We also denote vectors of the unknown parameters in (5) as $\phi = (a, b_r, c_i, \sigma_w^2, \sigma_{R^r}^2, \sigma_i^2)$ and the unknown parameters in (6) as $\psi = (\alpha_i, \beta_i)$ with $r = 1, 2, 3, 4, i = 1, 2, \dots, n$.

If we happened to know $\tilde{\xi}$, the state space system (5) - (6) collapses to a set of separate linear regressive equations (1)-(4) with known explaining variables, of which the procedure to estimate is well-known. We should be cautious, however, that the error term $\varepsilon_{i,t}$ in (1) is autocorrelated and can not be estimated directly. However, (1) could easily be transformed to usual regression equation with uncorrelated disturbance by simply multiplying both side of (1) with $(1 - c_i L)$, noting that (4) could be rewritten as $(1 - c_i L)\varepsilon_{i,t} = \eta_{i,t}$:

$$(1 - c_i L)y_{i,t} = \alpha_i(1 - c_i L)W_t + \beta_i(1 - c_i L)R_t^r + (1 - c_i L)\varepsilon_{i,t}$$

or

$$y_{i,t}^* = \alpha_i W_{i,t}^* + \beta_i R_{i,t}^{r*} + \eta_{i,t}, \quad \eta_{i,t} \sim N(0, \sigma_i^2) \quad (8)$$

where $y_{i,t}^* = (1 - c_i L)y_{i,t}$, $W_{i,t}^* = (1 - c_i L)W_t$ and $R_{i,t}^{r*} = (1 - c_i L)R_t^r$

Unfortunately, $\tilde{\xi}$ is unobserved. Traditionally, a state space system could be estimated by using Kalman filter to derive sample log likelihood conditional on the unknown parameters. The log likelihood is then maximized numerically with respect to these parameters until convergence. However, a weakness of this method is that log likelihood maximization is hard to perform when the number of unknown parameters becomes large. Stock and Watson (1998, 2001) propose a two-step procedure for state space model with very large cross-section dimension using principal components. Forni *et al* (2000) formulate an identification and estimation scheme for a generalized dynamic factor model, in which cross-section dimension goes to infinity. In this paper, we take the advantage of Bayesian Gibbs sampling procedure which allows us to easily estimate large cross-section state space system with large number of unknown parameters.

In Bayesian econometrics, unknown parameters are treated as random variables driven by underlying stochastic distributions. In our model, unknown parameter vectors ϕ , ψ and state vector $\tilde{\xi}$ are random variables whose mean and variance are to be estimated. Whereas traditional Kalman filter maximization infers $\tilde{\xi}$ conditional on the parameters, Bayesian inference on $\tilde{\xi}$ is based on joint distribution of $\tilde{\xi}$ and the unknown parameters ϕ and ψ . That is, we have to take

draws from the joint posterior distribution $p(\tilde{\xi}, \phi, \psi | \tilde{Y})$. Direct drawing from this joint distribution is impossible, however, since $p(\tilde{\xi}, \phi, \psi | \tilde{Y})$ does not take any well-known form. Nevertheless, the fact the state space system would collapse to estimable linear regressive equations when $\tilde{\xi}$ is known suggests that it is possible to take draws from the conditional density $p(\phi | \tilde{\xi}, \psi, \tilde{Y})$ and $p(\psi | \tilde{\xi}, \phi, \tilde{Y})$ using independent Normal Gamma prior. If we could take draws from the conditional density of $\tilde{\xi}$, $p(\tilde{\xi} | \phi, \psi, \tilde{Y})$, draws from the joint posterior distribution $p(\tilde{\xi}, \phi, \psi | \tilde{Y})$ could be derived using a Gibbs sampler.

The conditional density $p(\tilde{\xi} | \phi, \psi, \tilde{Y})$ could be obtained through a simulation smoother. Following Carter & Kohn (1994), $p(\tilde{\xi} | \phi, \psi, \tilde{Y})$ is given by (we hereafter omit ϕ, ψ in conditional part for ease of denotation).

$$p(\tilde{\xi} | \tilde{Y}) = p(\xi_T | \tilde{Y}) \prod_{t=1}^{T-1} p(\xi_t | \xi_{t+1}, \bar{y}_t), \text{ where } \bar{y}_t = (y_1, y_2, \dots, y_t)' \quad (9)$$

Because our model is Gaussian, the distribution of $\tilde{\xi}$ given \tilde{Y} is also Gaussian and could be written out more clearly as:

$$\xi_t | \bar{y}_t \sim N(\xi_{t|T}, P_{t|T}) \text{ or } \tilde{\xi} | \tilde{Y} \sim N(\tilde{\xi}_T, \tilde{P}_T) \text{ in stacked form} \quad (10)$$

for $t = T, T-1, T-2, \dots, 1$; $\xi_{t|T} = E(\xi_t | \bar{y}_t)$; $\tilde{\xi}_T = (\xi_{1|T}, \xi_{2|T}, \dots, \xi_{T-1|T}, \xi_{T|T})$;

$P_{t|T} = cov(\xi_t | \bar{y}_t)$ and $\tilde{P}_T = (P_{1|T}, P_{2|T}, \dots, P_{T-1|T}, P_{T|T})$.

We can directly compute $\xi_{t|T}$ and $P_{t|T}$ using recursive Kalman filter and smoother algorithm

with initial values $\xi_{1|0}$ and $P_{1|0}$:

$$\begin{aligned}
\xi_{t|t} &= \xi_{t|t-1} + P_{t|t-1}H'(HP_{t|t-1}H' + R)^{-1}(y_t - H\xi_{t|t-1}) \\
P_{t|t} &= P_{t|t-1} - P_{t|t-1}H'(HP_{t|t-1}H' + R)^{-1}HP_{t|t-1} \\
\xi_{t+1|t} &= F\xi_{t|t} \\
P_{t+1|t} &= FP_{t|t}F' + Q
\end{aligned} \tag{11}$$

and

$$\begin{aligned}
J_t &= P_{t|t}F'P_{t+1|t}^{-1} \\
\xi_{t|T} &= \hat{\xi}_{t|t} + J_t(\xi_{t+1|T} - \xi_{t+1|t}) \\
P_{t|T} &= P_{t|t} + J_t(P_{t+1|T} - P_{t+1|t})J_t'
\end{aligned} \tag{12}$$

The Kalman filter runs (11) forward with $t = 1, 2, \dots, T$ and (12) backward with $t = T - 1, T - 2, \dots, 1$. Details of the derivation of Kalman filter and smoother is presented in, for example, Hamilton (1994), chapter 13. Under the assumption that $\tilde{\xi}$ is stationary and the eigenvalues of F lie inside the unit circle, $\xi_{1|0}$ is zeros and $P_{1|0}$ is given by:

$$\text{vec}(P_{1|0}) = [I_{r^2} - (F \otimes F)]^{-1} \text{vec}(Q) \tag{13}$$

where \otimes denotes Kronecker product.

In principle, we could draw $\tilde{\xi}$ directly from the distribution in (10). However, Durbin and Koopman (2002) has developed a more efficient simulation smoother which facilitate the drawing

of $\tilde{\xi}$. In our model, their algorithm runs as follow:

First, draw a random vector v^+ from the density $p(v) = N(0, Q)$ in (7) and from v^+ recursively generate stacked vector $\tilde{\xi}^+$ and \tilde{Y}^+ using (5)-(6) with $\xi_{1|0}$ and $P_{1|0}$ computed from (13).

Second, compute $\tilde{\xi}_T = E(\tilde{\xi} | \tilde{Y})$ and $\tilde{\xi}_T^+ = E(\tilde{\xi}^+ | \tilde{Y}^+)$ using Kalman filter and smoother (11) and (12).

Finally, since $\tilde{\xi}_T^+ = E(\tilde{\xi}^+ | \tilde{Y}^+)$, $-\tilde{\xi}_T^+ + \tilde{\xi}^+$ gives \tilde{P}_T , where $\tilde{P}_T = cov(\tilde{\xi}^+ | \tilde{Y}^+)$. As \tilde{P}_T does not depend on data, $\tilde{P}_T = cov(\tilde{\xi}^+ | \tilde{Y}^+) = cov(\tilde{\xi} | \tilde{Y})$. Computing $\tilde{\xi} = \tilde{\xi}_T - \tilde{\xi}_T^+ + \tilde{\xi}^+$, we obtain a draw of $\tilde{\xi}$ from distribution $N(\tilde{\xi}_T, \tilde{P}_T)$.

Given the conditional distributions $p(\tilde{\xi} | \phi, \psi, \tilde{Y})$, $p(\phi | \tilde{\xi}, \psi, \tilde{Y})$ and $p(\psi | \tilde{\xi}, \phi, \tilde{Y})$, Gibbs sampling simulation is implemented to estimate $\tilde{\xi}, \phi$ and ψ from joint distribution density $p(\tilde{\xi}, \phi, \psi | \tilde{Y})$. Specifically, the Gibbs sampler would proceed in three steps:

1. Conditional on the parameter vectors ϕ and ψ , draw state vector $\tilde{\xi}$ from the conditional distribution $p(\tilde{\xi} | \phi, \psi, \tilde{Y})$ using Durbin and Koopman (2002) simulation smoother.
2. Conditional on the state vector $\tilde{\xi}$, draw parameter vector ϕ from the conditional distribution $p(\phi | \tilde{\xi}, \psi, \tilde{Y})$.
3. Conditional on the state vector $\tilde{\xi}$ and the parameter vector ϕ , draw parameter vector ψ from the conditional distribution $p(\psi | \tilde{\xi}, \phi, \tilde{Y})$.

Step 2 and 3 are carried out using independent Normal - Gamma priors.

These steps are iterated S times, of which the first S_o draws are discarded as burning-in replications to remove the effect of initial values. The initial values of ϕ and ψ are selected arbitrarily in the unit circle. To confirm simulation convergence, we compute Geweke (1992)

numeric standard error and Raftery and Lewis (1992) convergence diagnostics. We also repeat the Gibbs sampler with different initial parameter values to check if the effect of initial values is actually removed.

Two related identification problems should be noted when estimating the system, however. First, the signs of the common components and their associated coefficients in (1) are not separately identified. We handle this by requiring one of the coefficients for each component to be positive. Second, the scale of the those components and coefficients in are not separately identified either. We follow the convention to overcome this by normalizing the variances of η_t^w and η_t^r in (2)-(3) to unity.

The model is estimated using annual data from 10 East Asian countries, 12 European countries, three North America countries and nine South America countries over the period 1960 - 2002. GDP data is used as a proxy of aggregate output. For East Asian countries, we select Japan, Korea, China, HongKong, Taiwan and ASEAN 5 of Singapore, Thailand, Malaysia, Indonesia and the Philippines. For the purpose of our paper, we select only members of European Monetary Union (Austria, Belgium, Finland, France, Ireland, Italy, Luxemburg, Netherlands, Greece, Germany, Portugal and Spain) to present Europe. Three North America countries (NAFTA) are Canada, Mexico and the United States. Countries in South America group includes major members of the South American Community of Nations (Argentina, Bolivia, Brazil, Chile, Colombia, Paraguay, Peru, Uruguay and Venezuela). Data is drawn from World Bank World Development Indicator 2004 CD-ROM and IMF International Financial Statistics 2005 CD-ROM.

Before estimation, the data is filtered to generate detrended series. Although there are several detrending methods in the literature, we choose to use the most simple one: all the series are logged and first-differenced to achieve stationarity. We confirm the stationarity of these series

by applying ADF test for unit root. The tests failed to detect the presence of unit root in any series. To verify whether our result is robust with different filter scheme, we repeat our estimation with data filtered by another method, the Hodrick-Prescott (1980) filter. We take log of the data and apply this filter with parameter 100, keeping in mind that the Hodrick-Prescott filter might generate spurious correlation series (Harvey & Jaeger, 1993) and spurious cycles (Cogly & Nason, 1995). Detrended series are then standardized to obtain zero mean and unit variance. This is necessary to ensure that all series receive equal weights and every country is treated equally irrespective of its relative economic size. Estimation program is written in Matlab code using some modules from Jame LeSage's Econometric Toolbox.

III. Result:

The Gibbs sampler is iterated 12,000 times, of which the first 2,000 is discarded as burning-in replications. Geweke (1992) numeric standard error and Raftery and Lewis (1992) Z-test confirm Gibbs sampling convergence. Numeric standard errors are sufficiently small. Raftery and Lewis (1992) test (quantile 0.025, precision level 0.005 and associated probability of 0.95) indicates that the Gibbs sampler converges after less than 5000 replications. We also check if the initial values have worn off or not, using an informal procedure. Particularly, we repeat our simulation with random initial parameter values. Results are identical across repetition. Our estimation on Hodrick-Prescott filtered data and first-differenced data shows similarity in almost aspects. Specifically, both data produce similar dynamics of world and regional components, factor loading coefficients and their relative variance shares. While we base our assessment on first-differenced data, we will, however, refer to Hodrick - Prescott filtered data result wherever the discrepancy is significant and might lead to different interpretation. Estimated parameters are presented in Table 1 and Table 2. In Table 1, we show the autocorrelation coefficients of the world

and regional components estimated from equations (2) and (3). Factor loading coefficients and country-specific autocorrelation coefficients are shown in Table 2. The dynamics of world and regional components are illustrated in Figure 1, 2, 3 and 4.

Common component dynamics

Before investigating the relative role of the components in output volatility, it is necessary to take into account their dynamics. The purpose of this step is twofold. First, examining the dynamics of the components reveals whether the constructed common components do exhibit known cycles associated with major world and regional ups and downs in the past decades. If they appear to be poor indicators of world and regional business cycles, either our assumptions or estimation or both are inappropriate. Second, by analyzing the autocorrelation coefficients of the components, we are able to measure the persistence of world, regional and country-specific shocks. We interpret this persistence properties as measures of the speed of adjustment to shocks.

Figure 1 shows the dynamics of the world component, which could be interpreted as the world business cycle. Interestingly, the unobserved world factor well describe known economic events: the steady expansion of the early 1960s; the first steep fall of the mid 1970s due to the first oil price shock; the second big downturn of the early 1980s following the second oil price shock, the debt crisis in Latin America and the tight monetary policies in major industrial economies; the recession in the early 1990s and the recovery in the late 1990s, similar to those in Kose *et al* (2000) and Gregory *et al* (1997). The steepest drops in the figure coincide with the two severe recessions in mid 1970s and early 1980s. While Kose et at (2000) find that the recessions of mid 1970 and early 1980 are about of the same degree of severity, our world component indicates that the later is more disastrous.

<Figure 1 here>, <Figure 2 here>

<Figure 3 here>, <Figure 4 here>

East Asia regional factor is presented in Figure 2. Comparatively, the East Asia regional component seems to be less volatile except in 1997-1998 due to the East Asian financial crisis. Two other less severe output drops are found in mid 1970s and mid 1980s. The former might be reasonably explained by the first oil shock. The latter is associated with the regional recession in 1985-1986 in many of East Asia countries, such as HongKong, Singapore, Malaysia and the Philippines (in 1985, GDP growth rate of the Philippines fell to -7.6 percent (Lim and Bautista, 2002)). This is consistent with the common view that East Asia had experienced a long period of high and stable economic growth until the onset of the regional chaos (World Bank, 1993).

We depict the European regional component in Figure 3. The impact of first and second oil crises are clear in Europe. Unlike world common component, we find the first oil shock is more severe than the second to Europe. The early 1990s downturn associating with German unification and Exchange Rate Mechanism crisis was also more severe than the second oil price shock. We show North and South American regional components in Figure 4 and Figure 5. As we can see, North America cycle appears most volatile. The troughs and peaks of this cycle coincides remarkably with the US business cycles: recession in 1970, 1975, 1980, 1982, 1991, 2001 and expansion in 1973, 1981, 1984, late 1990s.

While the component dynamics reflect historical events, an investigation of their persistence property might provide insight of the adjustment speed. The persistence of the factors are measured by the first-order autocorrelation coefficients in Table 1. Larger coefficients mean higher degrees of persistence and imply longer effects of past shocks. Thus, we interpret the persistence property of the components as an indicator of adjustment speed.

The European regional component appears to be the most persistent. Its autocorrelation

coefficient is 0.51, indicating that adjustment to regional shocks in Europe is slow. North America responds fastest to regional shocks. Its autocorrelation coefficient is just 0.23. East Asia adjusts slightly slower, with the coefficient of 0.24 (if we use the Hodrick - Prescott filtered data, East Asia is even more responsive to regional shocks). South America is in the middle, with the coefficient of 0.36.

<Table 1 here>

Since a currency area involves a common monetary policy, regional adjustment speed is not influential on the cost of monetary policy abandonment. What matters is the responsiveness to asymmetric or country-specific shocks. In our model, these shocks are captured by country-specific components. Table 2 shows that European countries adjust fastest to country-specific shocks, followed by South America, North America and East Asia. European integration might be responsible for this rapid adjustment since factor movement, trade and investment within Europe probably alleviate adjustment process. Smaller countries often adjust faster to shocks. Adjustment speeds are fastest in the Netherlands, Portugal and Greece and slowest in Luxemburg, France and Ireland. In East Asia, asymmetric shocks have longest effects in Japan and the Philippines, reflecting structural and institutional rigidity in these countries. Korea responds fastest to country-specific shocks.

<Table 2 here>

Variance Decomposition

To measure the role of world, regional and country-specific shocks in output volatility, we conduct variance decomposition analysis. Under the assumption that the components are orthogonal, it is straightforward to decompose output variance into three parts corresponding to the three components, noting that $Var(\eta_t^w) = 1$ and $Var(\eta_t^r) = 1$ as assumed:

$$Var(y_i) = \alpha_i^2 Var(W) + \beta_i^2 Var(R_r) + Var(\varepsilon_i) \quad (14)$$

$$Var(W) = \frac{1}{1-a^2}; Var(R_r) = \frac{1}{1-b_r^2}; Var(\varepsilon_i) = \frac{\sigma_i^2}{1-c_i^2} \quad (15)$$

From (14) and (15), the share of output variance explained by those components are computed as

$$Var(y_i) = \frac{\alpha_i^2}{1-a^2} + \frac{\beta_i^2}{1-b_r^2} + \frac{\sigma_i^2}{1-c_i^2} \quad (16)$$

$$S_i^w = \frac{\frac{\alpha_i^2}{1-a^2}}{Var(y_i)}; S_i^r = \frac{\frac{\beta_i^2}{1-b_r^2}}{Var(y_i)}; S_i = \frac{\frac{\sigma_i^2}{1-c_i^2}}{Var(y_i)}$$

Where $r = 1, 2, 3$; $i = 1, 2, \dots, n$; S_i^w , S_i^r and S_i are the respective shares of world component, regional component and country-specific component in country i output variance.

The share of each component in output variance provides information on the symmetry of shocks, upon which one determines whether a country should join a regional currency area or not. If a large share of volatility in a country is explained by regional component, its shocks are synchronized with regional ones and the cost of forgoing independent monetary policy would be small. If the world component takes relatively large share, joining a broader monetary arrangement might be more appropriate. In case a large share of volatility is explained by country-specific variance, the country experiences asymmetric shocks and a regional currency area membership might be costly.

<Table 3 here>

Variance shares (in percentage) attributable to each component are presented in Table 3.

Surprisingly, the world component, on average, merely accounts for less than 10 percent of fluctuations in all regions. For many countries in East Asia such as Korea, China, and Thailand, the world shocks have virtually no impact. In the context of accelerating globalization, the finding that regional and country-specific factors are responsible for the most part of output variance is striking.

Country-specific factors account for a large share in output variance in all regions. They explain 63.5 percent and 52.3 percent of variance in East Asia and Europe respectively. For many countries, country-specific factor explains almost all output variance. In East Asia, Japan, the Philippines and China are most influenced by country-specific shocks. These factors account for at least a half of output variance in every country in East Asia. In Europe, the role of country-specific shocks is also significant in all countries, especially in Ireland where almost shocks are domestically spawned. Country-specific shocks are main source of volatility in South America. Only in the U.S., are country-specific shocks insignificant. This does not necessarily imply that shocks in the U.S. are mostly external. What it probably means is that most of shocks in the US spill over into global and regional shocks given the power of the U.S. economy.

Regional components explain significant shares of variance in all regions except South America. With no surprise, North America is the most synchronized region, specifically between the United States and Canada. In both countries, the regional factor explains more than two thirds of variance, probably due to strong integration between the economies. Mexico appears like an outlier as the regional factor accounts for just two percent of its output variance. Europe also appears synchronized with average regional share of 42 percent. Notably, the core countries of the EMU (Austria, Belgium, Germany, France, Italy and the Netherlands) exhibits strong common shocks, as shown by large shares (62%) of regionally rooted volatilities. The regional factor,

however, accounts for just a little output variance in Finland and Ireland; both are geographically far from the core. Regional factors are of little importance in South America.

In East Asia, the share of variance explained by regional factors is lower than in Europe. On average, regional factors accounts for 32 percent of output variance in East Asia. However, they have negligible impact on output variance in China and Japan. This is hardly surprising as both China and Japan have experienced quite different development paths. While China had undergone a centrally planned system before transitioning to market economy, Japan has been far more developed than other countries in the region. Quite strong synchronization is found between East Asian NICs, except for Taiwan. In Taiwan, world shocks and country-specific shocks are more important than regional shocks. In other NICs (Korea, HongKong, Singapore, Malaysia, Indonesia and Thailand), regional factors, on average, accounts for roughly 48 percent of output variance. The findings are consistent on both first-differenced data and Hodrick - Prescott filtered data.

Though East Asia shows certain degree of shock symmetry, it is hardly possible from the light of OCA criteria to say for sure whether East Asia is indeed qualified for a currency area. Fortunately, we can exploit the success of the European Monetary Union as a natural benchmark for comparison. If the role of regional factors in East Asia comes somewhat close to that in Europe, we might expect that it is feasible to replicate their success in East Asia. Otherwise, it might be difficult in the near future for such a plan to succeed. In general, East Asia exhibits less symmetry than Europe: variance share of regional component is lower and of country-specific component is higher, implying that East Asia is not as qualified as Europe for a currency area. Moreover, as shown above, adjustment to country-specific shocks in East Asia is slower and therefore, associating cost of adjustment is probably higher. However, the gap between East Asia and Europe

is not large (32% vs. 42%). If we compare the highly synchronized group of Korea, HongKong, Singapore, Malaysia, Indonesia and Thailand with European average, they appears to be potential candidates for a currency area. Obviously, this group cannot be compared with the European core where regional factor accounts for as large as 62 percent of output variance. It should be noted that Europe has achieved this level of synchronization after a long period of preparation and integration. Given the ongoing integration in East Asia, a vision of East Asia currency area would not be considered as utopian. Since not all East Asian countries are equally synchronized, a group of countries with higher degree of synchronization might form a currency union in the first place.

It is worthwhile to put our findings in juxtaposition. Bayoumi & Eichengreen (1994) find that East Asia is as plausible candidate as Europe for a currency area. They base their findings on correlation analysis of supply and demand disturbances identified by a structural VAR framework *à la* Blanchard & Quah (1989). Goto & Hamada (1994) and Goto (2002) find the same conclusion by exploring a principle component analysis. Sato *et al* (2003) extend the work of Bayoumi & Eichengreen (1994) using a three and five variables SVAR model. They find less persuasive support for a currency area in Asia and claim that only a subgroup of East Asian countries are possible candidates for monetary integration. They also find that adjustment speed to shocks is faster in East Asia. Chow & Kim (2003) modify the SVAR framework to identify shocks as global, regional and country-specific. Their result shows that in East Asia, country-specific shocks are more important and therefore, joining a currency area is not optimal. We find ourself in similar stance with Sato *et al* (2003) that, compared with Europe, East Asia is a less suitable but close candidate for a currency area. However, we distinguish between adjustment speeds to regional shocks and to country-specific shocks. We find the latter is slower in East Asia.

IV. Conclusion

In this paper, we examine the feasibility of forming a currency area in East Asia. Particularly, we check the symmetry of shocks by employing a Dynamic Factor Model to decompose aggregate output into three distinct components: world component, regional component and country-specific component. Our view is that, if output volatility in East Asia is dominantly driven by regional factors, then imposing a common currency would be desirable and entail little cost. On the contrary, if sources of fluctuations are largely country-specific, sacrificing monetary autonomy would be disastrous as a common policy is not sufficient to counter asymmetric shocks. In case world factors are important in determining output volatility, joining a broader currency arrangement might be more desirable.

Our result shows that East Asia is less plausible for a currency area than Europe in general. Furthermore, East Asian countries are less responsive to country-specific shocks so that adjustment process would be more costly and require longer time. However, a subgroup of East Asian countries including Korea, HongKong, Singapore, Malaysia, Indonesia and Thailand might be good candidates for a currency area since they appear highly synchronized. Our result is less optimistic than those of Bayoumi and Eichengreen (1994), Goto and Hamada (1994) and Goto (2002), where East Asia is as suitable as Europe for a currency area. However, we are not as pessimistic as Chow and Kim (2003), whose result shows that it is not possible for East Asia to form a currency area due to shock asymmetry. We find our result similar to Sato *et al* (2003). As Europe has experienced long integration and preparation process, a similar route might also bring East Asia closer to a currency union. A subgroup of more synchronized countries should go first, followed by other countries. Given East Asian accelerating integration in both trade and finance, it is not too early to prepare for such a move. Forming a currency area, however, requires not only economic condition but also political consensus, which might be more difficult to achieve in East

Asia.

There is still room for further improvement. First, since we work with output volatility in lieu of structural shocks, information on shocks could be conflated with policy responses. Our model could be extended to map the components' disturbances into structural shocks to avoid such conflation. A Factor-Augmented VAR model in the light of Bernanke *et al* (2005) and Stock and Watson 2005 might be the direction to take. Second, the model could also be extended to allow time-variant coefficients to capture structural changes. Finally, more aggregate variables, such as consumption, investment and price could be introduced into the model once data are available. We leave these for future works.

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Table 1: World and Regional Autocorrelation Coefficients

World	0.4724
East Asia	0.2415
Europe	0.5092
North America	0.2308
South America	0.3578

Table 2: Country Factor Loadings and Autocorrelation Coefficients

	World	Regional	Country-specific
East Asian			
Japan	0.20	0.19	0.66
Korea	0.02	0.68	0.19
China	0.01	0.00	0.31
Hongkong	0.31	0.51	0.49
Singapore	0.10	0.56	0.22
Malaysia	0.11	0.80	0.42
Indonesia	0.02	0.62	0.31
Philippines	0.06	0.37	0.56
Thailand	0.00	0.64	0.47
Taiwan	0.33	0.35	0.28
Average	0.12	0.47	0.39
Europe			
Germany	0.20	0.45	0.26
Belgium	0.21	0.57	-0.31
Finland	0.13	0.22	0.42
Neitherlands	0.19	0.56	0.09
France	0.17	0.57	0.34
Italy	0.19	0.50	0.29
Ireland	-0.05	-0.04	0.50
Spain	0.08	0.38	0.44
Portugal	0.20	0.51	0.04
Luxemburg	0.17	0.48	0.48
Austria	0.17	0.55	-0.26
Greece	0.21	0.35	-0.12
Average	0.16	0.42	0.18
North America			
US	0.15	0.89	0.32
Canada	0.16	0.74	0.48
Mexico	0.23	0.14	0.30
Average	0.18	0.59	0.36
Latin America			
Brazil	0.25	0.11	0.41
Argentina	0.22	0.18	0.05
Chile	0.14	0.14	0.31
Colombia	0.34	0.17	0.27
Peru	0.14	0.16	0.37
Uruguay	0.21	0.24	0.38
Paraguay	0.11	0.11	0.50
Venezuela	0.27	0.12	-0.04
Bolivia	0.07	0.05	0.26
Average	0.19	0.14	0.28

Table 3: Output Variance Decomposition

	World %	Regional %	Country-specific %
East Asian			
Japan	7.13	5.15	87.72
Korea	0.07	48.11	51.81
China	0.02	0.00	99.98
Hongkong	14.24	31.85	53.91
Singapore	1.36	37.84	60.80
Malaysia	1.44	67.15	31.41
Indonesia	0.06	45.92	54.02
Philippines	0.41	14.82	84.78
Thailand	0.00	54.99	45.01
Taiwan	17.92	16.02	66.05
Average	4.26	32.19	63.55
Europe			
Germany	15.73	83.10	1.17
Belgium	8.75	64.84	26.41
Finland	2.71	8.19	89.10
Neitherlands	6.28	55.52	38.21
France	5.89	71.52	22.60
Italy	6.16	46.35	47.50
Ireland	0.26	0.23	99.51
Spain	1.81	37.96	60.23
Portugal	6.19	42.60	51.20
Luxemburg	3.03	26.17	70.79
Austria	4.49	48.23	47.28
Greece	6.63	19.71	73.66
Average	5.66	42.03	52.30
North America			
US	3.32	95.66	1.02
Canada	4.24	77.13	18.63
Mexico	7.23	2.29	90.48
Average	4.93	58.36	36.71
Latin America			
Brazil	10.42	1.72	87.86
Argentina	9.08	5.49	85.43
Chile	2.54	2.47	95.00
Colombia	19.21	4.60	76.19
Peru	2.69	3.19	94.12
Uruguay	9.92	11.68	78.41
Paraguay	1.94	1.65	96.41
Venezuela	12.01	2.17	85.82
Bolivia	0.58	0.27	99.15
Average	7.60	3.69	88.71

Figure 1: World Common Component

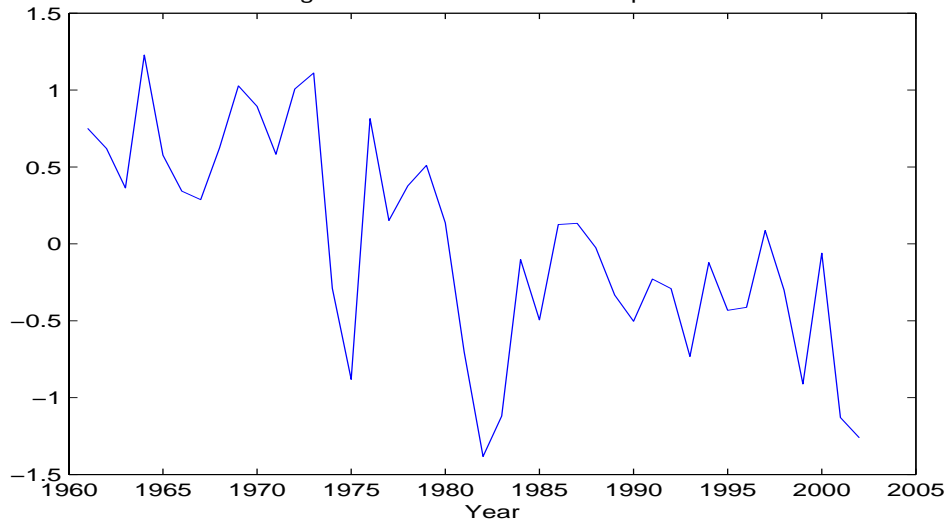


Figure 2: East Asian Regional Component

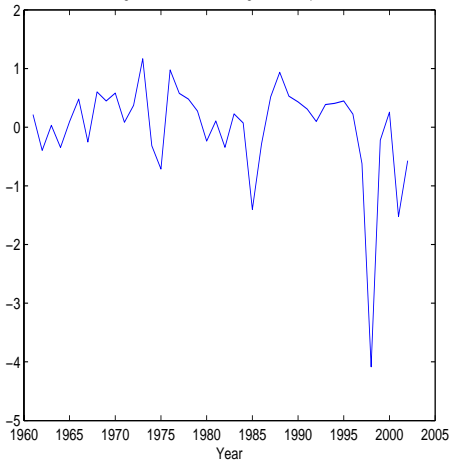


Figure 3: European Regional Component

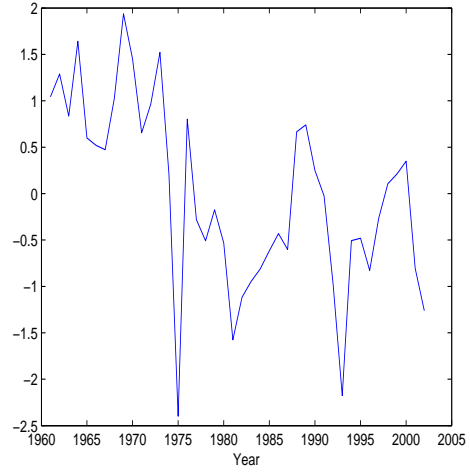


Figure 4: North American Regional Component

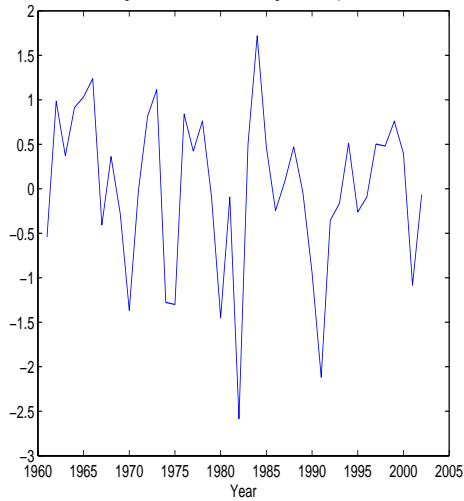


Figure 5: South American Regional Component

