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Measuring Intra-region Exchange Rate Variability and Its Path-through: the Case of ASEAN+3

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ABSTRACT

A multilateral currency union removes the intraregional exchange rates but not the union rate variability with the rest of the world. The intraregional exchange rate variability is thus latent. A two-step procedure is developed to measure the variability. The measured variables are used to model inflation and intraregional trade growths of individual union members. The resulting models form the base for counterfactual simulations of the union impact. Application to ASEAN+3 data shows that the intraregional variability consists of mainly short-run shocks, which have significantly affected the inflation and trade growths of the major ASEAN+3 members, and that a union would reduce inflation and promote intraregional trade at the union level but the benefits facing each member vary and may not be significant enough to warrant a vote for the union.

Key words: currency union, latent variables, dynamic factor model, simulation

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1. Introduction

This research is motivated by the growing attention on evaluating the cost and benefit of having a currency union. Policy interest and debates over the union have undoubtedly been kept high by the launch of the Eurozone in 2002 and the subsequent expansion of the EU. But the concerns extend well beyond Europe. For instance, monetary union is on the political table in East Asia, where the US dollar is used effectively as a common currency anchor and the competitive devaluation strategy has been widely adopted in the wake of the Asian financial crisis (e.g. see McKinnon and Schnabl, 2004). The sustainability of the anchor is now under question as continued devaluation of the US dollar challenges many economies, especially those which have been experiencing a buoyant growth and mounting foreign currency reserves.

Should countries of a region choose currency union to mitigate foreign exchange risk and uncertainty? This question has challenged economists for decades. Theoretical work on the optimal currency areas goes back to the seminal paper by Mundell (1961). More recent contributions include Alesina and Barro (2002), Ca'Zorzi *et al* (2005) and Sanchez (2005).[1] A notable feature of the recent models is the increasingly explicit treatment on the impact of the exchange rate variability in terms of shocks or risks. Ca'Zorzi *et al* (2005) demonstrate that it is the variance of real exchange rate shocks with respect to the currency union, rather than the deterministic factors of the rate, which plays a vital role in measuring the expected loss function of the union. Further decomposition of such shocks at a country level is proposed by Sanchez (2005) on the basis of the observation that the exchange rate dynamics may differ considerably among the individual members of a perspective union.

On the empirical front, results vary due mainly to differences in the choice of modelling methods, data selection and data processing. There is generally a lack of empirical studies testing those shock-based postulates from the recent theoretical models. The lack, we believe, is caused by a number of difficulties in measuring and identifying the exchange rate variability that a union is to remove. For example, what variables should be used to represent the variability? Should they correspond to the nominal or the real exchange rate changes? Barro and Tenreyro (2007) use the standard deviation of the residuals from a second-order autoregression of bilateral real exchange rates (BREER) between two economies as a measure of price-comovement shocks; Tenreyro (2007) defines the variability variable by the annual standard deviations of monthly nominal exchange rate changes. It seems that both types should be considered as they could impact on an economy in different manners. Once the variability is represented by nominal and real exchange rate shocks, a more serious problem arises. How can we identify from the shocks the exchange rate variability of a perspective union, which is not due to the variability of the rest of the world? Surely, a union only removes the intraregional exchange rates but not its rate variability with the world. The problem has not been explicitly dealt with in the literature as far as we know. Most of the relevant studies examine bilateral country data, corresponding to the theoretical framework of a bilateral union, e.g. in Alesina *et al* (2002). However, the phenomenon of variable co-varying a great deal with factors outside the bilateral environment is found prominent in these studies. Tenreyro (2007) treats the phenomenon as an endogeneity problem and proposes to tackle it by the instrument variable (IV) method (see also Barro and Tenreyro, 2007). It is however difficult to see how the method can work at the country level in the context of a multilateral currency union.

[1] Ca'Zorzi *et al* (2005) also give a relatively detailed literature survey; see Frankel (2004) and Frenkel and Nickel (2005) for more literature coverage.

Apart from the measurement problems, model choice poses another contentious issue. One popular choice is the ‘gravity equation’ model for trade among the international economist circle (e.g. see Tenreyro, 2007). But the absence of dynamic specification makes it ill-suited for evaluating the impact of the exchange rate variability. The use of dynamic panel models or even individual dynamic models at the country level seems a more promising alternative, such as the studies on exchange rate pass-through by Gagnon and Ihrig (2004), Artis and Ehrmann (2006), Bussière (2007).

This study attempts to resolve three issues. First, we maintain that the intraregional exchange rate variability is latent and measurable by means of dynamic factor models (DFM); a two-step procedure is designed for filtering out the part of the exchange rate variability with the rest of the world. Second, we believe that the transmission paths of the purified regional variability to individual countries of the region should differ and be modelled by dynamically adequately specified models; the general→specific dynamic modelling approach is adopted in our country-level modelling and one of the advantages of the approach is that it enables us to categorise the purified regional factors into two types: short-run shocks and long-run shocks due to purchasing power parity (PPP) misalignments. Third, we differentiate the situation of the significant presence of the explanatory variables of interest from the situation whether such variables would significantly alter the explained variables of interest; we find it practically inadequate to deduce economic relevance from the former situation alone, as commonly reported in most of the academic literature,[2] and therefore employ the counterfactual simulation/forecasting method to examine the latter situation.

We apply our new approach to the case of ASEAN+3, using monthly time-series data for the period of 1990M1-2007M9. Briefly, our results show that the intraregional exchange rate variability of ASEAN+3 consists of mainly short-run shocks, that the magnitude of the shocks has remained undiminished in the post-crisis period, that the shocks exert significantly impact on the inflation and trade growths of the major ASEAN+3 members, and that a union would reduce inflation and promote intraregional trade at the union level but the benefits enjoyed by each member vary and may not be significant enough to warrant a vote for the union.

The rest of the paper is organised as follows. The next section describes the new method; section 3 reports the main application results to ASEAN+3, and the last section concludes with the main findings.

2. Methodology

We take PPP as our theoretical base of measurement. Following the convention, denote the bilateral real exchange rates (BREER) of one country vis-à-vis one foreign country at time t as:

$$(1)\qquad r_t = \left(\frac{e_f p}{p_f}\right)_t$$

where e_f is the nominal exchange rate denominated in the currency of the foreign country, p and p_f are price indices of the domestic and foreign countries respectively. The bilateral PPP framework is often extended to a multilateral setting when the empirical interest is in certain regional issues with panel data sets at hand. Denoting the set of all countries by $N = \{1,2,\cdots,n\}$, the set of foreign countries vis-à-vis country i, the domestic country of

[2] The problem was raised and criticised by McCloskey (1985) though little heed has been paid to the criticism.

interest, by $N_{-i} = \{1, 2, \cdots, i-1, i+1, \cdots n\}$. In view of a perspective currency union, N is further divided into two subsets – one for the union region (superscripted by u) and the other the rest of the world (superscripted by w): $N = \{N^u, \quad N^w\}$. To bridge this multilateral setting with model (1), we choose to represent the notional foreign entity vis-à-vis country i by latent common factors extracted from N_{-i}. The choice is made from the consideration that, while high degrees of correlation exist, the idiosyncratic features in the price and the exchange rate data of individual countries are too significant to sustain the idealised conditions assumed by the theory underlying (1). The gap between data and theory is in effect a measurement error problem when country-level panel data are used to estimate PPP-based models, see Qin *et al* (2007) and Qin (2008) for more discussion. Here, two sets of common factors F_{it}^u and F_{it}^w, can be extracted for each country $i \in N^u$ corresponding the two geographical regions by means of the DFM:

$$(2) \qquad \begin{aligned} (Z_{if})_t^j &= \Gamma_{if}^j F_{it}^j + \varepsilon_{if,t}^j \\ F_{it}^j &= \Lambda_i^j(L) F_{it-1}^j + v_{it}^j \end{aligned} \qquad j = u, w$$

where $(Z_{if})_t^j$ denotes a vector of the relevant foreign-country related variables, $f \in N_{-i}$, F_{it}^j denotes a vector of latent common factors whose dimension is a lot smaller than that of $(Z_{if})_t^j$.

Obviously, we cannot use F_{it}^u to represent the intraregional exchange rate variability for country i if there is significant correlation between F_{it}^u and F_{it}^w. To filter out the correlation, we adopt the following simple system of equations:

$$(3) \qquad F_{it}^u = \Pi_i F_{it}^w + \Phi_{it}^u$$

where Φ_{it}^u is a vector of the 'purified' regional factors representing the intraregional exchange rate variability.[3]

Let us now turn to the issue of modelling the transmission paths of the purified regional variability to individual countries. Under the PPP, the dynamics of a panel of BREER with respect to the partner foreign countries has been modelled by a VAR (Vector AutoRegression) allowing for unit roots, e.g. in Koedijk *et al* (2004):

$$(4) \qquad \ln(r_{if})_t = \alpha_i + \beta_i \ln(r_{if})_{t-1} + \gamma_i(L)(\dot{r}_{if})_{t-1} + u_{it}$$

where $(\dot{r}_{if})_t = \ln(r_{jf})_t - \ln(r_{if})_{t-1}$ denotes growth rate, $\gamma_i(L)$ is a finite-order lag polynomial and u_{it} is white-noise residual. Notice that (4) is effectively what underlies the augmented Dickey-Fuller test for unit roots. For the purpose of studying the exchange rate path-through to domestic inflation, we relax the homogeneous restriction implied by $\gamma_i(L)$ on the dynamic adjustments speeds of the domestic price and the foreign price in (4) and reformulate it into an error-correction model (ECM) using (1):

$$(5) \qquad \dot{p}_{it} = \alpha_i(L)\dot{p}_{it-1} + \alpha_f(L)(\dot{p}_f - \dot{e}_f)_t + \beta_i \ln(r_{if})_{t-1} + u_{it}$$

where $\alpha_i(L)$ and $\alpha_f(L)$ are finite-order lag polynomials, and β_i is the feedback coefficient with the implicit condition of $\beta_i < 0$ under the long-run PPP hypothesis. Notice that (5) decomposes the dynamic movement of BREER into four types of shocks – short-run

[3] Lee *et al* (2004) choose to filter out the world's impact from the regional output and trade by means of a multi-component factor model. Their approach amounts to filtering F_i^u and F_i^w together in one DFM, which would require more degrees of freedom in practice.

domestic inflation, short-run foreign inflation in combination with exchange rate variability, long-run disequilibrium due to PPP misalignment and a residual term. Since the second and the third types embody exchange rate variability from the external sources, we thereafter refer to them as the short-run and the long-run foreign shocks respectively. Notice also that the first three types of shocks are not only structurally interpretable but also relatively uncorrelated.

In the event that the foreign shocks are latent but measurable by DFMs, we can modify (5) into a dynamic factor error-correction model (DF-ECM). For example, the DF-ECM of inflation for country *i* with respect to the impact of the regional exchange rate variability of a perspective union region can be written as:

(6) $$\dot{p}_{it} = \alpha_i(L)\dot{p}_{it-1} + A_f(L)\Phi\left(\dot{p}_f - \dot{e}_f\right)_t^u + B_i\Phi\left(r_{if}\right)_{t-1}^u + u_{it}^u \qquad f \in N_{-i}{}^u, N^u{}_{-i} = N_{-i} \cap N^u$$

Equation (6) shows that we need two types of foreign shocks for country *i*, i.e. the long-run and the short-run shocks. Hence, we set up two types of indicator sets, which are further divided into the union region and the rest of the world:

Long-run indicators: $\begin{cases} \ln\left(r_{if}\right)_t^u, & f \in N^u{}_{-i} \\ \ln\left(r_{if}\right)_t^w, & f \in N^w \end{cases}$; short-run indicators: $\begin{cases} \left(\dot{p}_f - \dot{e}_f\right)_t^u, & f \in N^u{}_{-i} \\ \left(\dot{p}_f - \dot{e}_f\right)_t^w, & f \in N^w \end{cases}$.

Using each type as $\left(Z_{if}\right)_t^j$ in turn and running (2) and (3), we obtain a set of long-run factors, $\Phi\left(r_{if}\right)_t^u$, and a set of short-run factors, $\Phi\left(\dot{p}_f - \dot{e}_f\right)_t^u$, for each member country.

Equation (6) also tells us that currency unification does not guarantee inflation reduction. The effect of exchange rate variability via the long-run factors can be inflation-stabilising through disequilibrium-correction, whereas the effect via the short-run factors depends on the signs of $A_f(L)$.[4] It is thus an empirical matter to assess whether the overall effect of a currency union on inflation is positive or negative. The assessment can be done via a counterfactual model simulation exercise, once (6) is estimated. The exercise involves running ex-post dynamic forecasting, for the sample period, of the following DF-ECM:

(7) $$\hat{\dot{p}}_{it} = \hat{\alpha}_i(L)\dot{p}_{it-1} + \hat{A}_f(L)\Phi\left(\dot{p}_f\right)_t^u + \hat{B}_i\Phi\left(\frac{p_i}{p_f}\right)_{t-1}^u$$

where the initial values of the lagged $\dot{p}_{it-j}$ are assumed zero, the coefficients denoted with 'hat' are those estimated from (6), and $\Phi\left(\dot{p}_f\right)_t^u$ and $\Phi\left(\frac{p_i}{p_f}\right)_t^u$ are obtained by a two-step procedure: (a) Re-run the DFMs (2) for the union case using two indicator sets net of the exchange rates; (b) using the resulting two sets of common factors to filter $\Phi\left(\dot{p}_f - \dot{e}_f\right)_t^u$ and $\Phi\left(r_{if}\right)_t^u$ respectively via (3). A comparison of $\{\dot{p}_{it}\}$ and $\{\hat{\dot{p}}_{it}\}$ will tell us whether the union is inflation mitigating or amplifying.

A more challenging issue is to evaluate the impact of a currency union on other macro variables of the union members. Such evaluations entail macroeconometric model simulations (see Hughes *et al*, 1992), and are beyond the scope of this study. However, it is

[4] Hughes *et al* (1992) give more detailed economic argument on the possible negative effect of a currency union; they argue that the union may incur the loss of the absorptive mechanism of exchange rate variability among its members. Artis and Ehrmann (2006) use structural VAR models to examine the shock-absorbing and shock-generating effects of exchange rates.

feasible to estimate how much the removal of the intraregional exchange rate shocks would benefit the intraregional trade. The estimation can be based on a DF-ECM extension of the 'gravity equation' model for trade:

$$(8) \qquad \Lambda_{0,i}\begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_t^u = \Lambda_i(L)\begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_{t-1}^u + \Gamma_f(L)\Phi\left(\dot{e}_f\right)_t^u + \left\langle \Pi_i(L)\dot{\delta}_t + B_i\begin{pmatrix}\Phi_x(*)\\ \Phi_m(*)\end{pmatrix}_{t-1}\right\rangle + \begin{bmatrix}\varepsilon_x\\ \varepsilon_m\end{bmatrix}_{it}$$

where $\dot{x}$ and $\dot{m}$ are the intraregional export growth and import growth respectively, Λ_0 is a non-diagonal coefficient matrix, $\dot{\delta}$ represents other short-run variables and $\Phi_x(*)$ and $\Phi_m(*)$ are the long-run error-correction factors. One difficulty with (8) is the identification of those long-run factors. In the commonly used gravity equation models (see Eaton and Kortum, 2001 and Anderson and Wincoop, 2003), the components of the long-run factors include GDP and geographic distance of the trading partners involved. Empirically, trade flow variables are often found to be cointegrated with GDP and/or other real variables. Nominal variables, such as prices or exchange rates, are only present in terms of short-run shocks. That feature enables us to exploit the property of a growth model in that its coefficient estimates should be unbiased when the level variables are non-stationary and cointegrated with other non-stationary variables in the long run (see Hendry, 1995; Chapter 7). In other words, the estimation of the impact of $\Phi\left(\dot{e}_f\right)_t^u$ is achievable by truncating (8) into a growth model:

$$(9) \qquad \Lambda_{0,i}\begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_t^u = \Lambda_i(L)\begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_{t-1}^u + \Gamma_f(L)\Phi\left(\dot{e}_f\right)_t^u + \begin{pmatrix}u_x\\ u_m\end{pmatrix}_{it}$$

where $\Phi\left(\dot{e}_f\right)_t^u$ is the complement of $\Phi\left(\dot{p}_f\right)_t^u$ in (7). Similar to (7), the trade impact of a currency union can be assessed by running ex-post dynamic forecasting on:

$$(10) \qquad \hat{\Lambda}_{0,i}\begin{bmatrix}\hat{\dot{x}}_i\\ \hat{\dot{m}}_i\end{bmatrix}_t^u = \begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_t^u - \left\langle \hat{\Lambda}_i(L)\begin{bmatrix}\dot{x}_i\\ \dot{m}_i\end{bmatrix}_{t-1}^u + \hat{\Gamma}_f(L)\Phi\left(\dot{e}_f\right)_t^u\right\rangle$$

Again, the initial values of the endogenous variables are assumed zero, and the coefficients denoted with 'hat' are those estimated from (9).

Methodologically, our approach bears close similarity to the method of latent variable structural equation models (e.g. see Wansbeek and Meijer, 2000). An illustration of our approach for the case of the inflation model is given by the path diagram in Figure 1. As seen from the diagram, the intraregional exchange rate variability is represented by latent variables and DFM (2) is used as the primary measurement model for these variables, whereas model (6) is the structural model.

Before moving on to the empirical application, a number of technical issues need to be briefly explained concerning the implementation of our method (see also Qin, 2008 for more description). First, two recently developed procedures of consistent estimators are used to determine the number of factors used in (2). One is developed by Bai and Ng (2007) and the other by Onatski (2005). The larger of the two estimates is adopted when they differ. Next, the Kalman filter algorithm is used for extracting the factors of (2), with the initial parameter estimates obtained via principal component analysis, as developed by Camba-Mendez *et al* (2001). As for equation (3), the estimation is carried out by OLS (Ordinary Least-Squares). OLS is the principle method used for models (6) and (9) as well, since the latent regressors there effectively play the role of 'instrumental variables' in correcting for measurement error attenuation. The computer-automated model reduction software PcGets (see Hendry and Krolzig, 2001), is employed for primary model simplification search for (6) and (9), or 'testimation' using the software's terminology, because the number of parameters tends to be quite large. The key advantage of PcGets is that it carries out testimation by the general → specific approach in a consistent and efficient manner such that the specific model resulted from testimation is guaranteed to be data-coherent and parsimoniously encompassing of the general model at the starting point (see Hendry, 1995; Chapter 9). Notice that (9) assumes simultaneity between $\dot{x}_t$ and $\dot{m}_t$. Accordingly, 2SLS (Two-Stage Least-Squares) is used during the testimation. As the dynamic forecasting with (7) and (10) assumes estimated parameter constancy from (6) and (9), the specific models resulted from PcGets testimation are further verified, and sometimes revised using PcGive, taking advantage of the special facility of the software on checking parameter constancy via recursive estimation (see Hendry and Doornik, 2001).

3. Application: the case of ASEAN+3

This section presents the empirical results of applying the method described above to the region of ASEAN+3. ASEAN is currently composed of Brunei, Cambodia, Indonesia, Laos, Myanmar, Malaysia, Philippines, Singapore, Thailand and Vietnam.[5] The addition of Japan, South Korea and China defines ASEAN+3. The general economic situation of ASEAN+3 in relation to a perspective currency union has been studied recently by Lee *et al* (2004), Sanchez (2005) and Zhang and Yin (2005). Ito and Sato (2006) and Cortinhas (2007) estimate the exchange rate pass-through of the major ASEAN countries via VAR models.

In the present investigation, Hong Kong and Taiwan are also included in the regional data set, making our ASEAN+3 region cover fifteen economies, i.e. $n^u = 15$. However, the country-level analysis is only carried out for eight ASEAN+3 countries, namely, ASEAN-5 – Indonesia, Malaysia, Philippines, Singapore and Thailand, and the 'plus three' – China, Japan, and South Korea. Figure 2 plots the ratios of the intra ASEAN+3 trade to the total trade of these eight economies by country. To form the region defining the rest of the world, twenty six countries are selected, i.e. $n^w = 26$, including India, Pakistan and the rest being OECD countries other than Japan and Korea, namely Austria, Australia, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, Turkey, UK, and USA. These two geographic sets are simply referred to as the regional set and the world set respectively in the subsequent text.

Monthly time-series data are collected for the period of 1990M1—2007M9. Since there was an East Asian financial crisis in the third quarter of 1997, the modelling exercise is carried out for two sub-sample periods: the pre-crisis period: 1990M1—1997M6, and the post-crisis period: 1998M10—2007M9. The data coverage includes consumer price indices (CPI),[6] dollar denominated exchange rates, bilateral exports and imports of the countries within ASEAN+3. Due to data constraints, the trade data set covers only for the post-crisis period. Hence, model (9) is estimated for the post-crisis period. Detailed information about all the data series is given in the Appendix.

[5] Detailed information on ASEAN is available at: http://www.aseansec.org/.

[6] Trade price indices are not used here, partly because of data unavailability and partly because of the very different compositions of the traded goods underlying the indices among the countries concerned here.

3.1 Measurement of Regional Shocks

For each of the eight ASEAN+3 countries, two sets of common factors are extracted from the two geographic indicator sets respectively. Each set is comprised of two types of shocks: the long-run shocks and the short-run shocks, each type corresponding to one run of DFM (2). The indicator sets are standardised as required of factor models. The factor extraction is carried out twice time-wise: once for the pre-crisis period and the second time for the post-crisis sample period.[7] The numbers of factors estimated for each type of the shocks range between 4-6 for the individual countries, as shown in Table 5 and Table 6. The lag lengths are chosen by reference to the Akaike and Schwarz information criteria. One lag is found to be adequate for most cases; a few go up to three lags. Once all the factor sets are extracted, the purified ASEAN+3 shocks are obtained by regressing the regional factors on the world factors using (3).

To evaluate how much the regional factors co-vary with the world factors, redundancy analysis is carried out between F_{it}^{u} and F_{it}^{w} of the various types of shocks. As shown from the redundancy statistics in Table 1, a large part of the variance in the long-run regional factors is explained by the world factors, especially for the period prior to the Asian crisis; the degree of dependency has dropped considerably in the post-crisis period. Meanwhile, the within-group correlation of the purified long-run shocks among the eight countries remains high, as shown from the canonical correlations between the first factors in Table 3, although there is a slight decrease in the correlations of the '+3' countries in the post-crisis period. In contrast, the regional short-run factors, $F\left(\dot{p}_f - \dot{e}_f\right)_t^u$, are much more independent of the world factors (see Table 1), implying that the purified $\Phi\left(\dot{p}_f - \dot{e}_f\right)_t^u$ retain most of $F\left(\dot{p}_f - \dot{e}_f\right)_t^u$; the independence seems to stem mostly from inflation volatility as higher dependency is found between the short-run factors purely for the exchange rates, $F\left(\dot{e}_f\right)_t^u$ and $F\left(\dot{e}_f\right)_t^w$, as shown in Table 1. Notably, a majority of these short-run factors are more independent of the world factors for the post-crisis period, reflecting the fact that most of the countries adopted cautious monetary policies in the wake of the East Asian crisis. Further cross-country comparison of the purified short-run shocks tells us that, except for the China case of the pre-crisis period, the volatility of these shocks is similar (expressed in standard deviations in Table 2), and that all these shocks remain highly correlated (expressed in canonical correlation between the first pair of canonical variables in Table 4). There is no evidence that the shocks facing smaller members are larger than those facing larger members. The similarity is actually easily explained by the fact that the regional short-run indicator sets only differ by one indicator between different countries while they share the same world indicator set. The above results suggest that the latent regional exchange rate variability stem mainly from short-run shocks exposing commonly to individual members. Reduction of the variability would, therefore, call for a currency union. But the issue of whether the individual members have enough incentive to vote for the union would largely depend on the severity of the impact that the variability impinges on their economies.

[7] Due to insufficient degrees of freedom for the pre-crisis period, the pre-crisis factor set is obtained by truncating the factors obtained through running (5) for the full-sample period.

3.2 Modelling the Transmission Paths of Regional Shocks

To assess the severity of the impact, the purified shocks are used to explain inflation and the intraregional trade growths of the eight countries respectively via models (6) and (9). Table 5 and Table 6 report the summary results of the PcGets testimation on (6) and (9). As seen from these tables, the testimation results in a much smaller specific model than the starting general model with nine lags for each country. For example, the regressors in (6) for the China case of the post-crisis period are reduced from 48 to 6, as shown in the note of Table 5.

In the case of modelling inflation by (6), the short-run regional shocks remain to be significantly present in all the country cases for both the prior and post crisis periods, whereas the long-run shocks are present in half of the countries for the pre-crisis period and reduced to three for the post-crisis period. Comparing the long-run results with the finding by Qin *et al* (2007) that the long-run common factors are all significantly present in the inflation equations of the relevant Asian countries, it indicates that, for the countries under investigation, much of the disequilibrium-correcting mechanism concerning real exchange rate adjustments is with respect to the world rather than the ASEAN+3 region. The cost of a currency union is rather low as far as the potential loss in the intraregional real rate shock-dampening capacity is concerned. The impact that the regional rate variability impinges on individual countries is predominantly by the short-run shocks. The transmission paths of the impact differ significantly across countries and across samples as well, as can be seen from the different lag patterns shown in Table 5.[8] The change of the lag patterns between the pre-crisis and post-crisis results may reflect the policy regime shift in the exchange rate system of the countries concerned in the wake of the crisis.

As for the results of the intraregional trade model (9), the short-run exchange rate shocks remain to be significantly present, see Table 6. Noticeably, the lag patterns in the trade cases are more complicated than those in inflation cases, especially on the longer lags. This finding confirms the common sense that prices normally react more instantly than trade to exchange rate variability. Interestingly, our trade model results contradict Tenreyro's finding (2007) that nominal exchange rate variability has no effect on trade. We believe that Tenreyro's finding is the result of an imprecise measure of the exchange rate variability compound with inadequately specified dynamic models.

As pointed out in Section 2, the significant presence of the short-run shocks in both the inflation and trade models does not necessarily imply that the removal of these shocks would reduce inflation and/or promote trade by significant amounts. To examine that issue, ex post dynamic forecasts are carried out using (7) and (10). The means and variances of the simulated counterfactual forecasts are then compared to those of the actual sample series. Tests of statistical differences are also calculated. As shown from the summary results in Table 7, a perspective currency union would reduce the inflation level of the whole region by 3-6%, increase intraregional export and import growths by 16% and 10% respectively, and dampen the volatility in the trade growths as well, though it would only dampen the inflation volatility for the pre-crisis period. These results corroborate Rose's (2004) finding that there is relatively strong evidence of trade promotion of a currency union. On the whole, the scheme is found to be union-wide beneficial, though the benefits are not so significantly certain if judged by the test statistics.

Let us now look at the country-level results in Table 8 and Table 9. Among the 'plus three' countries, China would benefit from both a reduction of inflation coupled and an increase in the regional trade growths; Japan would enjoy import growth and a slight

[8] Detailed estimation results at the country level are not reported in order to keep the paper short.

reduction in the volatilities of inflation and the export growth as well; the benefits for Korea include inflation reduction, export growth and reduction in trade growth volatility. However, all these benefits are associated with relatively high uncertainty, except for the reduction of import growth volatility in the Korean case, when the significance test statistics are considered. As for the ASEAN-5 countries, Malaysia appears to gain relatively the most benefits, which include reduction in the volatilities of inflation and import growth, as well as higher trade growth; the benefits for Singapore lie mainly in the volatility reduction of both inflation and trade growths; the results for the three countries are rather mixed; there is certain volatility reduction in the trade growths in the Indonesian case; the only visible benefit for the Philippines is some volatility reduction of inflation and export growth, whereas what Thailand might enjoy is inflation reduction according to the post-crisis sample result. Again, few of these benefits are warranted with significance test statistics, similar to the 'plus three' cases. Interestingly, our results share certain similarities with the findings by Cortinhas (2007), though the modelling methods differ.

Ca'Zorzi *et al* (2005) show in their theoretical model that economically smaller countries would gain more than relatively larger countries in a multilateral currency union. Comparing the benefits enjoyed by the '+3' group as versus those by the ASEAN-5 group, we see no evidence in support of their theory. Alesina and Barro (2002) maintain that countries with higher inflation volatility would benefit more from the union than those with lower volatility. Our empirical results provide little evidence for their theory. The lack of theory-expected evidence is mainly the result of the substantial difference found in the short-run shock transmission paths among different member countries. The difference largely reflects the fact that there exists great difference in the history of the foreign exchange and trade policy arrangements among these countries. The smaller countries or those with more volatile inflation history may well have been more open to the world economy than those larger economies or those with more stable inflation record, as certainly the case of ASEAN+3.

4. Conclusions

The present study is motivated by a number of problems revealed in the empirical studies of the impact of a multilateral currency union. A new approach is proposed for the empirical evaluation of the impact of the union for its member countries at the country level. By exploiting DFMs, we are able to obtain a common measure of the intraregional exchange rate variability which a country of the region faces specifically. The measure is 'purified' not only of those idiosyncratic shocks from individual countries within the region, but also of the impact from the world common exchange rate variability outside the region. The measure is further categorised into two types: short-run and long-run common currency shocks. These shocks are used as explanatory variables to model the inflation and intraregional trade growths of the country concerned. The resulting models provide us with a base to simulate and evaluate the counterfactual situation of how much inflation and trade growths would be affected by the removal of these shocks. Methodologically, our approach can be considered as a special case of the latent variable structural models used commonly in behavioural research.

Application of our approach to eight major ASEAN+3 countries yields a number of interesting findings. First of all, the regional long-run exchange rate variability covariates with the world exchange rate variability a great deal whereas the short-run exchange rate variability is mainly regional specific. Consequently, a currency union would result in reducing the intraregional short-run currency volatility risks without much loss of the regional capacity of assimilating disequilibrium risks from the world currency movement.

Moreover, our dynamic modelling results show that the regional short-run shocks exert significant impact on the inflation and the intraregional trade growths of all the countries studied, overshadowing the impact found of the regional long-run shocks. We also find that the dynamic transmission paths of the regional shocks differ significantly from country to country. The finding makes it an oversimplified statement that smaller countries would benefit more than larger countries from a currency union. The benefit of a currency union is found, however, to be less substantial as far as the model-simulated magnitudes in inflation reduction and trade promotion are concerned. At the regional level, the magnitudes in trade promotion are much larger than the amount of inflation being reduced; at the country level, results vary and, in many cases, the benefits may not to be considered as substantial enough to warrant a vote for the union.

Appendix: Variables and data sources

Variable	Economy	Source	Particulars
CPI	Brunei, Cambodia, Indonesia, Laos, Myanmar, Malaysia, Philippines, Singapore, Vietnam, Japan, Korea, China, Hong Kong, India, Pakistan, Austria, Australia, Belgium, Canada, Denmark, Finland, France, Greece, Germany, Ireland, Italy, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, Thailand, Turkey, UK, USA	IMF *International Financial Statistics* (code I64)	Australia: quarterly series; New Zealand: monthly data to 1992M12; the later series is derived from quarterly data; China: data prior to 1993 are from the State Bureau of Statistics; Brunei, Laos, Myanmar, Vietnam: shorter than full-sample series
	Taiwan	Datastream	
	Czech Republic	Datastream	Data start from 1993M1
US$ exchange rate	As above	Datastream	Czech Republic, Vietnam: shorter than full-sample series
Export to ASEAN+3 Import from ASEAN+3	China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Thailand	IMF *Direction of Trade Statistics*	Aggregate series from bilateral trade series of these countries vis-à-vis individual ASEAN+3 economies
	Taiwan	Datastream	Bilateral trade series vis-à-vis ASEAN-5 and 'plus three'
Export to world Import from world	China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Thailand	IMF *Direction of Trade Statistics*	

Note: All the series are monthly for the period of 1990M1 — 2007M09 except for those noted in the particulars.

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Figure 1. Path Diagram of Modelling Inflation by Latent Variables of the Regional Exchange Rate Variability

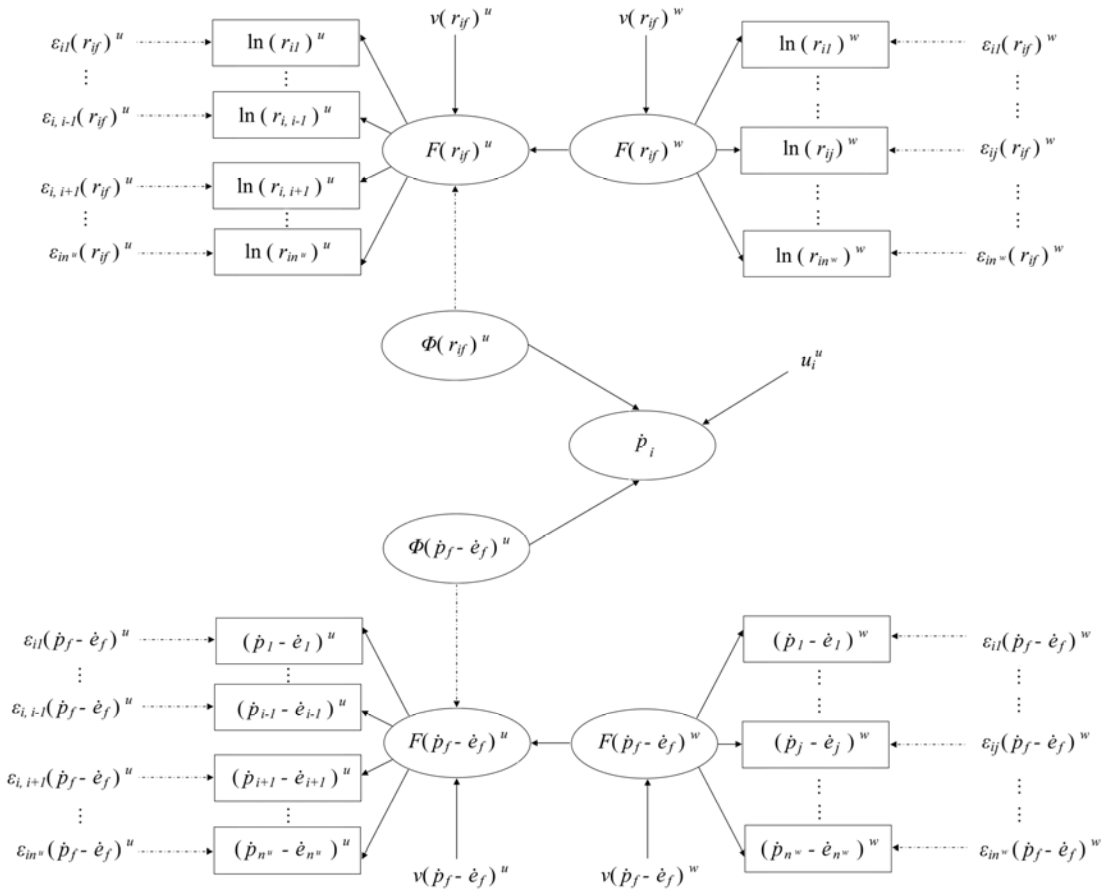
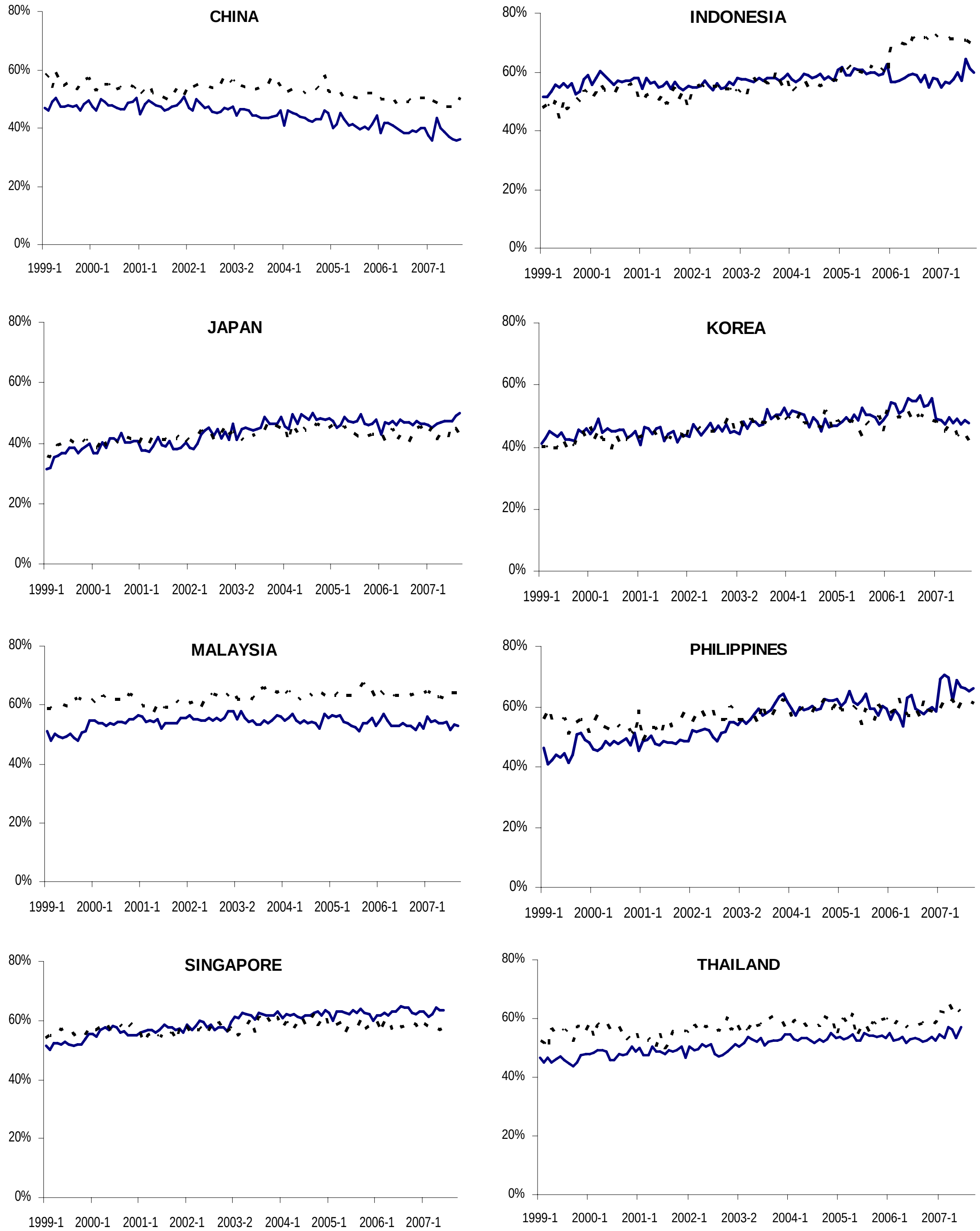


Figure 2. Ratios of Exports and Imports Within ASEAN+3 to the Total Exports and Imports (Solid line: exports; dotted line: imports)



Note: The data series are derived from IMF *Direction of Trade Statistics*, see Appendix.

Table 1. Redundancy Statistics of the Regional Factors as Explained by the World Factors

	Long run: $F(r_{if})_t^u$		Short Run: $F(\dot{p}_f - \dot{e}_f)_t^u$		Short Run: $F(\dot{e}_f)_t^u$	
	Sample 1	Sample 2	Sample 1	Sample 2	Sample 1	Sample 2
Indonesia	0.8469	0.5147	0.1777	0.0968	0.5352	0.3251
Malaysia	0.8160	0.6124	0.1573	0.1100	0.5167	0.4419
Philippines	0.8569	0.5850	0.1251	0.0991	0.5447	0.4396
Singapore	0.8699	0.5370	0.1436	0.0814	0.4780	0.4944
Thailand	0.7973	0.6351	0.1691	0.0899	0.5329	0.4813
China	0.8381	0.6099	0.0419	0.1025	0.5313	0.5127
Japan	0.9544	0.5874	0.1264	0.0933	0.5014	0.5009
Korea	0.8343	0.5009	0.1820	0.0887	0.5580	0.4196

Note: Sample 1 covers the pre-crisis period of 1990M01-1997M06; sample 2 covers the post-crisis period of 1998M10-2007M09.

Table 2. Standard Deviations of the Purified Shocks

	Long run: $\Phi(r_{if})_t^u$		Short Run: $\Phi(\dot{p}_f - \dot{e}_f)_t^u$		Short Run: $\Phi(\dot{e}_f)_t^u$	
	Sample 1	Sample 2	Sample 1	Sample 2	Sample 1	Sample 2
Indonesia	1.1911	4.7444	1.0077	0.7928	0.6901	0.4886
Malaysia	1.5804	2.7124	1.0189	0.8888	0.6827	0.5842
Philippines	1.7837	1.4325	1.0284	0.7930	0.7273	0.5747
Singapore	2.1404	1.9779	0.9688	0.7900	0.6341	0.6252
Thailand	1.7838	1.7456	1.0070	0.7952	0.6955	0.5735
China	1.7967	1.6819	1.9487	0.8277	1.7921	0.6127
Japan	1.2276	1.7355	1.1125	0.8138	0.7373	0.6063
Korea	1.5482	0.9520	1.0066	0.9013	0.7119	0.5643

Note: Sample 1 covers 1990M01-1997M06; sample 2 covers the post-crisis period of 1998M10-2007M09. The factors are pooled into one series in the calculation.

Table 3. Pair-wise Canonical Correlation of the Purified Regional Long-run Shocks

		Sample 2							
		China	Indonesia	Japan	Korea	Malaysia	Philippines	Singapore	Thailand
Sample 1	China		0.8393	0.9727	0.8880	0.9008	0.9322	0.9475	0.9175
	Indonesia	0.9900		0.9011	0.9191	0.9838	0.9701	0.9792	0.9935
	Japan	0.9661	0.9459		0.9433	0.8887	0.9299	0.8753	0.8671
	Korea	0.9648	0.9848	0.9495		0.8307	0.9261	0.8711	0.9609
	Malaysia	0.9439	0.9854	0.9505	0.9866		0.9601	0.9856	0.9825
	Philippines	0.9551	0.9766	0.9525	0.9708	0.9708		0.9590	0.9997
	Singapore	0.9673	0.9550	0.9761	0.9696	0.8976	0.9413		0.9715
	Thailand	0.9592	0.9845	0.9814	0.9843	0.9843	0.9931	0.9579	

Note: Sample 1 covers 1990M01-1997M06; sample 2 covers the post-crisis period of 1998M10-2007M09. The statistics reported here are the first canonical coefficients.

Table 4. Pair-wise Canonical Correlation of the Purified Regional Short-run Shocks for the Post-crisis Period

		$\Phi\left(\dot{p}_f - \dot{e}_f\right)_t^u$							
		China	Indonesia	Japan	Korea	Malaysia	Philippines	Singapore	Thailand
$\Phi\left(\dot{e}_f\right)_t^u$	China		0.9979	0.9992	0.9980	0.9978	0.9985	0.9981	0.9986
	Indonesia	0.9993		0.9955	0.9965	0.9946	0.9975	0.9967	0.9970
	Japan	0.9989	0.9984		0.9980	0.9980	0.9964	0.9993	0.9993
	Korea	0.9998	0.9990	0.9988		0.9998	0.9952	0.9970	0.9929
	Malaysia	0.9996	0.9983	0.9983	0.9999		0.9947	0.9939	0.9936
	Philippines	0.9996	0.9993	0.9988	0.9996	0.9995		0.9986	0.9996
	Singapore	0.9963	0.9931	0.9975	0.9944	0.9910	0.9962		0.9994
	Thailand	0.9997	0.9996	0.9979	0.9983	0.9995	0.9989	0.9985	

Note: The post-crisis period covers 1998M10-2007M09. The statistics reported here are the first canonical coefficients.

Table 5. Impact of Regional Factors on Inflation by Model (6)

Samples	Pre-Asian crisis			Post-Asian crisis		
ASEAN+3	Number of long-run factors	Number of short-run factors	Short-run factor lag structure	Number of long-run factors	Number of short-run factors	Short-run factor lag structure
Indonesia	$4 \to 0$	$4\times9 \to 4$	2, 5, 6	$4 \to 0$	$4\times9 \to 4$	2, 3, 5, 6
Malaysia	$6 \to 2$	$4\times9 \to 10$	0, 1, 2, 3, 4, 5, 6	$4 \to 0$	$6\times9 \to 4$	0, 1, 3
Philippines	$5 \to 0$	$4\times9 \to 2$	0, 2	$5 \to 2$	$4\times9 \to 6$	2, 4, 5
Singapore	$4 \to 0$	$4\times9 \to 6$	0, 1, 4, 6	$4 \to 0$	$4\times9 \to 2$	2, 6
Thailand	$6 \to 3$	$4\times9 \to 8$	1, 2, 3, 5, 6	$5 \to 2$	$4\times9 \to 4$	0, 1, 7
China	$4 \to 2$	$4\times9 \to 7$	1, 2, 6	$4 \to 1$	$4\times9 \to 4$	0, 2, 3
Japan	$4 \to 0$	$4\times9 \to 2$	4, 6	$4 \to 0$	$4\times9 \to 7$	0, 3, 5, 7
Korea	$6 \to 2$	$4\times9 \to 5$	2, 3, 5, 6	$4 \to 0$	$6\times9 \to 11$	1, 2, 3, 4, 5

Note: The column 'number of long-run factors' shows the number of long-run factors extracted using DFM (2) and also how many of them remaining in the specific model through PcGets model reduction. The column 'number of long-run factors' shows similar information, with an additional part showing the maximum lag used in the general model at the starting point of the PcGets model reduction; nine lags are used here. The column 'short-run factor lag structure' lists the lags of the short-run factors remaining in the specific model at the final stage of model reduction. The following gives an example of the China inflation model for the post-Asian crisis period:

$$\hat{\dot{p}}_t = 0.0117 - 0.49\,\dot{p}_{t-3} - 0.0055\,\phi(\dot{p}_f)_{2,t} - 0.0024\,\phi(\dot{p}_f)_{2,t-2}$$
$$+ 0.0024\,\phi(\dot{p}_f)_{3,t-2} - 0.0021\,\phi(\dot{p}_f)_{3,t-3} + 0.0012\,\phi(r^*)_{1,t-1} + f(seasonals)$$

Table 6. Impact of Short-run Regional Factors on Trade by Model (9**)**

Trade	Exports			Imports		
ASEAN+3	Simultaneity with imports	Number of short-run factors	Short-run factor lag structure	Simultaneity with exports	Number of short-run factors	Short-run factor lag structure
Indonesia	no	$4\times9 \to 9$	5, 6, 8, 9	no	$4\times9 \to 7$	5, 6, 7, 8
Malaysia	no	$4\times9 \to 7$	3, 4, 5, 7, 8	no	$6\times9 \to 15$	1, 2, 3, 4, 5, 6, 7, 8, 9
Philippines	no	$4\times9 \to 14$	1, 2, 3, 8, 9	no	$4\times9 \to 5$	5, 6
Singapore	no	$4\times9 \to 10$	1, 2, 3, 4, 8, 9	no	$4\times9 \to 6$	1, 2, 4, 6, 8, 9
Thailand	no	$4\times9 \to 14$	0, 1, 4, 5, 8, 9	no	$4\times9 \to 9$	0, 2, 4, 6, 7
China	yes	$4\times9 \to 4$	2, 5, 8	yes	$4\times9 \to 6$	2, 5, 6, 7, 8
Japan	yes	$4\times9 \to 13$	0, 2, 3, 4, 6, 8, 9	no	$4\times9 \to 8$	0, 1, 2, 5, 8
Korea	no	$4\times9 \to 9$	0, 1, 3, 5, 6, 8, 9	no	$6\times9 \to 27$	0, 1, 2, 4, 5, 6, 7, 8, 9

Note: The columns on 'simultaneity' show whether current export or import variable is present in the specific model through PcGets model reduction. When 'yes', two-stage least squares (2SLS) method is used for model estimation. For the columns 'number of short-run factors' and 'short-run factor lag structure', see the note in Table 5 above for the detailed explanation. The sample covers the post-crisis period of 1998M10-2007M09.

Table 7. Union-wide Effects

Inflation (monthly)					
1990M8 – 1997M6			1999M10 – 2007M9		
Average / standard dev.	Simulated as % of actual	Significance test statistics	Average / standard dev.	Simulated as % of actual	Significance test statistics
0.0048	96.95%	$z = 0.2667$	0.0024	93.9%	$z = 0.4315$
0.0107	89.55%	$F = 1.25$***	0.0067	100.4%	$F = 0.9909$
Export growth (monthly)			Import growth (monthly)		
1999M10 – 2007M9			1999M10 – 2007M9		
0.0106	116.7%	$z = -0.349$	0.0114	109.95%	$z = -0.235$
0.1002	98.36%	$F = 1.0337$	0.0959	96.75%	$F = 1.0682$

Note: 'dev.' stands for deviation; *** indicates significance at 1%.

Table 8. Country-level Effects: Plus-three (the results are based on monthly rates)

	Average / standard dev.	Simulated as % of actual	Significance test statistics	Average / standard dev.	Simulated as % of actual	Significance test statistics
China		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0077	71.6%	$z = 1.0201$	0.0012	56.03%	$z = 0.4165$
	0.0139	97.1%	$F = 1.0603$	0.0088	102.7%	$F = 0.9475$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0172	107.7%	$z = -0.071$	0.0167	119.9%	$z = -0.143$
growth	0.131	103.1%	$F = 0.9399$	0.1273	98.6%	$F = 1.0283$
Export	0.0169	114.4%	$z = -0.12$	0.0172	148.1%	$z = -0.315$
growth	0.139	106.8%	$F = 0.8761$	0.1432	101%	$F = 0.9796$
Japan		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0011	134.8%	$z = -0.59$	-0.0002	91.1%	$z = -0.052$
	0.0042	99.46%	$F = 1.0109$	0.0029	96.7%	$F = 1.0695$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0077	111.9%	$z = -0.069$	0.0091	109.6%	$z = -0.049$
growth	0.0936	101.6%	$F = 0.9686$	0.0958	103.1%	$F = 0.9411$
Export	0.0089	113.2%	$z = -0.071$	0.011	80.1%	$z = 0.107$
growth	0.1197	95.9%	$F = 1.087$	0.1145	95.5%	$F = 1.0963$
Korea		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0045	96.8%	$z = 0.2121$	0.0024	94.9%	$z = 0.2001$
	0.0044	95.7%	$F = 1.0923$	0.0042	106.2%	$F = 0.8854$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.012	98%	$z = 0.027$	0.0122	100.5%	$z = -0.005$
growth	0.0703	81.4%	$F = 1.51$**	0.0726	83.2%	$F = 1.443$*
Export	0.0105	118.6%	$z = -0.205$	0.0134	110.1%	$z = -0.116$
growth	0.0694	94.99%	$F = 1.108$	0.0656	94.86%	$F = 1.1113$

Note: 'dev.' stands for deviation; * and ** indicate significance at 10% and 5% respectively.

Table 9. Country-level Effects: ASEAN-Five (the results are based on monthly rates)

	Average / standard dev.	Simulated as % of actual	Significance test statistics	Average / standard dev.	Simulated as % of actual	Significance test statistics
Indonesia		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0065	89.99%	$z = 0.6237$	0.0064	98.9%	$z = 0.0497$
	0.0064	107.1%	$F = 0.8719$	0.01	99.5%	$F = 1.0095$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0193	101.6%	$z = -0.023$	0.0242	94.4%	$z = 0.0734$
growth	0.101	91.1%	$F = 1.205$	0.11	84.6%	$F = 1.397$*
Export	0.0115	116.3%	$z = -0.201$	0.0147	96.3%	$z = 0.0456$
growth	0.0682	93.04%	$F = 1.155$	0.0667	93.65%	$F = 1.14$
Malaysia		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.003	101.6%	$z = -0.101$	0.0016	101.5%	$z = -0.068$
	0.031	99.95%	$F = 1.0009$	0.0025	97.48%	$F = 1.0525$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0072	107.8%	$z = -0.044$	0.0064	119.4%	$z = -0.068$
growth	0.094	93.77%	$F = 1.137$	0.1005	98.4%	$F = 1.034$
Export	0.0086	105.3%	$z = -0.035$	0.0097	114.5%	$z = -0.079$
growth	0.0914	102.5%	$F = 0.9513$	0.0902	114.8%	$F = 0.7593$
Philippines		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0074	92.87%	$z = 0.4894$	0.0035	102%	$z = -0.095$
	0.007	98.09%	$F = 1.0393$	0.0054	93.6%	$F = 1.1425$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0088	93.57%	$z = 0.0431$	0.0124	83.2%	$z = 0.1263$
growth	0.0925	102.1%	$F = 0.9597$	0.09	99.7%	$F = 1.006$
Export	0.0134	89.3%	$z = 0.100$	0.015	67.2%	$z = 0.2502$
growth	0.1097	84.9%	$F = 1.389$*	0.1135	90.2%	$F = 1.228$
Singapore		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.0047	139.8%	$z = -0.568$	0.0015	90.2%	$z = 0.1152$
	0.0228	84.4%	$F = 1.405$*	0.0091	97.7%	$F = 1.0478$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0083	103.8%	$z = -0.029$	0.0122	81.2%	$z = 0.1556$
growth	0.0783	97.4%	$F = 1.0542$	0.0812	98.1%	$F = 1.0396$
Export	0.0116	94.6%	$z = 0.0479$	0.0157	74.9%	$z = 0.228$
growth	0.0952	96.3%	$F = 1.078$	0.0964	96.3%	$F = 1.079$
Thailand		1990M8 – 1997M6			1999M2 – 2007M9	
Inflation	0.004	105.2%	$z = -0.282$	0.0019	82.8%	$z = 0.5853$
	0.0047	104%	$F = 0.9241$	0.004	99.5%	$F = 1.0101$
		1999M5 – 2007M9			2002M10 – 2007M9	
Import	0.0123	111.9%	$z = -0.119$	0.0144	103.8%	$z = -0.033$
growth	0.0894	95.7%	$F = 1.0927$	0.0896	102.2%	$F = 0.9576$
Export	0.0128	91.9%	$z = 0.0852$	0.0151	88.8%	$z = 0.1063$
growth	0.0841	105.5%	$F = 0.8984$	0.0853	103.7%	$F = 0.93$

Note: 'dev.' stands for deviation; * and ** indicate significance at 10% and 5% respectively.