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**The Effects of Productivity Gains in Asian Emerging Economies:  
A Global Perspective**

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# The Effects of Productivity Gains in Asian Emerging Economies: A Global Perspective.

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## Abstract

This paper investigates international responses of key macroeconomic variables, particularly real exchange rates, to simultaneous shocks to productivity in the traded sector in eight Asian emerging and developing countries. We use panel estimation techniques to construct component submodels in a thirty country global vector autoregressive (GVAR) model. The GVAR approach can account for interaction among all countries and capture many potential international transmission channels. We identify the shocks by using sign restricted impulse responses. We find that increases in traded-sector productivity in Asian developing countries lead to a real appreciation of the domestic currencies, in line with the Balassa-Samuelson hypothesis. Inflation also increases in many Asian developing countries. After the shocks, nontraded sector productivity in the US and other developed countries increases, suggesting that there is a compositional shift in their production, away from the traded goods toward the nontraded goods. This allows productivity in the nontraded sector to increase. Further, the traded sector productivity shocks in Asia stimulate international trade in most countries.

## 1 Introduction

Emerging market economies in Asia have been growing rapidly in recent decades, transitioning to free-market-oriented economies and gradually integrating with global markets. This economic success was once known as the "Asian Miracle". More recently, large economies such as China and India have become major economic forces within Asia and worldwide, given their rapid economic growth, their expanding markets and their roles in international markets via trade, capital flow and investment. The main drivers of economic growth in these emerging countries include the reallocation of factors to higher productivity sectors, and rising global trade reflecting trade liberalization and expanding supply chains (Cubeddu et al (2014)). However, although emerging Asia's economic dynamism has had increasing international significance, there are few studies that investigate the effects of rapid economic growth and productivity gains in these countries on the economic well-being of the US, Europe, Australia and the rest of the world.

This paper undertakes an analysis in a multi-country framework to assess the dynamic movement of key economic variables in response to simultaneous shocks to traded-sector productivity improvement in eight Asian emerging and developing economies including China, India, Indonesia, Thailand, Philippines,

Malaysia, Pakistan and Sri Lanka, and to examine how the shocks spill across the rest of the world. We place an emphasis on how real exchange rates respond to the shocks, and investigate the importance of traded sector productivity gains in explaining real exchange rate movements in each country.

It is widely accepted that the Balassa-Samuelson (BS) hypothesis (Balassa 1964, Samuelson 1964) is a plausible theory which explains the long-run relationship between sectoral productivity differentials and real exchange rates, so we consider it as a starting point for our analysis. The BS hypothesis is that countries with faster productivity growth in traded sectors relative to nontraded sectors will tend to experience a real exchange rate appreciation. As productivity growth in traded sectors is normally linked to economic growth, this leads to a prediction among economists that the real exchange rates of emerging economies in Asia will appreciate against other slow growing economies. However, the validity of this prediction is arguable, as empirical evidence on the BS hypothesis is mixed. Further, there has been recent international political pressure on fast growing Asian countries (e.g. China) to appreciate their currencies.

The estimation of a global model using an unrestricted vector autoregressive (VAR) model is infeasible in practice so we turn to the global vector autoregressive (GVAR) model, originally introduced by Pesaran et al. (2004) and developed by Dees et al. (2007b) to construct a truly international framework for the study of real exchange rate behaviour. Compared to the existing literature, the use of a GVAR model is more appealing for this study because the GVAR approach can account for interaction between a large number of countries and capture many potential international transmission channels. This is a crucial feature, given that real exchange rate movement of one particular country is affected by not only a change in the domestic economy, but also by changes in many other economies through international markets. Therefore, ignoring inter-country linkages can cause misleading results. Also, the inclusion of many countries in the model allows us to answer questions that cannot be tackled in previous studies.

We construct a multi-country model composed of thirty countries. This group of countries include eight Asian emerging and developing countries and another twenty-two countries that represent other developed and developing countries around the globe. Our data set is a panel of annual time series with different time spans starting from 1970 and extending to 2008, covering real exchange rates, main exchange rate fundamentals (productivity differentials between traded and nontraded sectors, relative price of nontraded goods to traded goods, real GDP, government consumption shares, terms of trade and international trade shares), inflation, short-term nominal interest rates and oil prices. The individual country models contain domestic variables, country-specific foreign variables and a common global variable (i.e. the oil price). The country-specific foreign variables are constructed as country-specific trade weighted averages of the

corresponding domestic variables for all other countries or regions. The country models are constructed separately in a way that allows for the possibility of common trends among variables across countries. The key assumption for estimation and inference is that foreign variables and a global common variable are weakly exogenous, compatible with a limited degree of weak dependence across idiosyncratic shocks. The link matrices, which are defined as country-specific trade weights, allow individual country variables to be transformed into global variables. Then individual country models are combined together via the link matrices to complete the GVAR model.

Although existing empirical literature on the relationship between sectoral productivity differentials and real exchange rates is large, it is limited to a group of developed countries and it typically uses a blunt proxy of traded and nontraded productivity differentials such as per capita GDP. We construct sectoral productivity differentials and relative prices of nontraded goods to traded goods by using a newly developed traded-nontraded industrial sector classification (Dumrongittikul, 2012) which explicitly considers sector-based productivity and price series for each industry. We add to the empirical literature on the behaviour of the real exchange rate in a multi-country setting by including both developing and developed countries around the world. The literature related to this field has remained sparse due to the data limitations associated with the developing countries and the difficulties of dealing with the "curse of dimensionality". Our study is different from the previous GVAR literature in two important ways. First, most previous studies have employed generalized impulse response functions (GIRFs), proposed by Koop et al. (1996) and developed by Pesaran and Shin (1998) to examine the international effects of shocks. GIRFs can deal with the cross correlation of shocks within each country and across different countries, but they make economic interpretation difficult because they are associated with a set of non-orthogonalized shocks. We avoid this problem by using sign restrictions on impulse response functions to identify the shocks. This latter approach requires only a set of a priori plausible restrictions suggested by economic theory, and the global model offers an easy and straight-forward way to incorporate a large number of reasonable sign restrictions that can be used for identifying the shocks that are of interest. Second, to the best of our knowledge, this is the first work in the GVAR literature that constructs country-group models using panel estimation techniques, instead of estimating individual country models separately, in order to alleviate estimation issues that arise from limited data and to increase the efficiency of the parameter estimates.

Our main finding is that real exchange rate responses to traded sector productivity shocks seem to follow the BS prediction. That is, productivity gains in the traded sector relative to the nontraded sector in Asian developing countries cause their real exchange rates to appreciate in the medium run, and also lead to

short run appreciations in most non-US developed and other developing countries. The real value of the US dollar declines. Interestingly, we find that positive shocks to traded-nontraded productivity differentials in Asia also lead to a decline in traded-nontraded productivity differentials for the US and non-European developed countries in the medium run. This implies increases in productivity in the nontraded sector relative to the traded sector in these developed countries, suggesting that when Asian developing countries become more productive in traded goods, the developed countries shift their production and resources away from traded sectors toward nontraded sectors. Moreover, the shocks can lead to either positive or negative effects on real GDP in foreign economies, i.e. short-run increases in real GDP in Europe and many developing countries, and medium-run declines in US, Japanese and Australian GDP. Gains in traded sector productivity stimulate international trade among countries as shown from increased imports and exports in most countries, corresponding to our expectation. In addition, variance decomposition analysis shows that fluctuations in real exchange rates explained by the productivity shocks are varied and quite limited in all countries examined.

The remainder of the paper is organized as follows. We briefly present the literature most relevant to our study in Section 2. We provide details on the data in Section 3. In Section 4, we set out the country/region-specific models and outline the way that we combine the country/region-specific models into the GVAR framework. Section 5 provides the details for our identification schemes, persistence profiles, impulse responses and forecast error variance decomposition. Section 6 reports the empirical results and Section 7 concludes.

## 2 Review of the Literature

Much of the literature on the relationship between productivity growth and real exchange rates assesses the BS hypothesis. However, empirical studies of the BS hypothesis have provided conflicting evidence. Bergin et al. (2006) and Lee and Tang (2007) find a positive association between these two variables, whereas this relationship is weak in studies by Drine and Rault (2005), and Peltonen and Sager (2009). As argued in the literature, the puzzling findings might be a consequence of using a weak proxy of productivity such as per capita GDP, manufacturing output per employee or real value added per employee, even if it is appropriately disaggregated into sectors (see Drine and Rault (2005), Chong et al. (2012) for example). In addition, as the BS hypothesis relies on several underlying assumptions including the Law of One Price in traded goods, frictionless intra-economy labour mobility, perfect capital markets and slow productivity growth in nontraded sectors, the invalidity of one or more of these assumptions might undermine the prediction of the BS hypothesis.

In terms of productivity shock identification, recent research has employed the sign restriction methodology. This approach has become increasingly popular over the recent years because conventional identification approaches require many restrictions, most of which seem to be too stringent. However, most of this work aims to investigate the effects of productivity on labour markets. Dedola and Neri (2007) consider the dynamic effects on hours worked, inflation and the short-term interest rate in the US. Similarly, Peersman and Straub (2009) focus on hours worked and employment in the Euro Area. There are only a few studies on the real exchange rate reaction to productivity shocks that use a sign restricted impulse response approach. Corsetti et al. (2009) use sign-identified VAR models of the US to identify a shock to tradable goods productivity in the US. They account for international transmission with the use of all the variables in the form of cross-country differentials between the US and an aggregate of industrialized countries. Similarly, Enders et al. (2011) estimate a VAR model of the US and they derive sign restrictions based on a quantitative general equilibrium model using a wide range of plausible model parameterisations. They find similar results to Corsetti et al. (2009), in that productivity gains cause a short-run increase in the relative price of domestically produced goods and US currency appreciation in the short run. However, Enders et al. (2011) find that in the medium run the shock leads to a currency depreciation.

Hiebert and Vansteenkiste (2007)'s study is possibly the most closely related to our paper, in that they investigate the effects of technology shock in the GVAR framework. Their GVAR is constructed by combining 12 industry-specific models of US manufacturing, allowing for an assessment of spillovers from industry-specific shocks to manufacturing subsectors. However, while our study examines responses of macro variables in different countries to traded-sector productivity shocks, they focus on the impacts of a common technology shock (proxied by US manufacturing sector spending on research and development) on US manufacturing labour markets. Further, they investigate the dynamics of the transmission of shocks using generalised impulse response rather than response that has been identified using sign restrictions.

In short, the existing literature on identifying traded sector productivity shocks has relied on two-country models ("Home" and "Foreign") and has focused on industrialized countries. In this paper, we provide new evidence on responses to simultaneous shocks to productivity in traded sectors relative to nontraded sectors in eight Asian emerging and developing economies (China, India, Indonesia, Thailand, Philippines, Malaysia, Pakistan, and Sri Lanka). We have paid careful attention to the construction of the traded-nontraded productivity variable by using the method introduced by Dumrongritikul (2012) for classifying industries. To our knowledge, the study of the dynamic responses of macroeconomic and financial variables to traded-sector productivity shocks has not been explored yet in a global framework.

### 3 Data

We employ the entire data set used in Dumrongrittikul (2011) and Dumrongrittikul (2012). The data set covers thirty countries and includes annual time series with different time dimensions from 1970 to 2008. Following Pesaran et al. (2004), we treat five countries (France, Germany, Italy, Netherlands and Spain) that were original members of the European monetary union in 1999 as a single economy. In particular, domestic variables of this region are constructed as a weighted average of corresponding country-specific series using the weights based on a ratio of country's PPP-adjusted GDP to the total PPP-adjusted GDP of these five countries. See Table B1 for the list of countries.

We consider the following variables in the model: the logarithm of the real effective exchange rate ( $q_{i,t}$ ); the inflation rate ( $\pi_{i,t} = p_{i,t} - p_{i,t-1}$  where  $p_{i,t}$  is the consumer price index); the logarithm of the ratio of traded to nontraded productivity measures ( $x_{i,t}$ ); the logarithm of real GDP ( $y_{i,t}$ ); the logarithm of the ratio of government consumption to GDP ( $gov_{i,t}$ ); the logarithm of the openness of the economy ( $open_{i,t}$ ), measured as the ratio of the sum of exports and imports to GDP; the nominal interest rate ( $si_{i,t} = \ln(1 + NI_{i,t}/100)$  where  $NI_{i,t}$  is the short-term interest rate per annum measured as a percentage); the log oil price index ( $oil_t$ ). The sources for the data are documented in Dumrongrittikul (2011) and Dumrongrittikul (2012).

### 4 Global VAR Modelling

We consider  $N + 1$  countries or regions, indexed by  $i = 0, 1, 2, \dots, N$ , in the global model. Due to the dominant role of the US in the world economy, the US is treated as the reference country and labelled country 0. Throughout this analysis, we use a superscript \* for foreign variables and no superscript for domestic variables (except the oil price,  $oil_t$ ). All variables are in log form. Let  $g_{i,t}$  be a  $k_i \times 1$  vector of domestic variables, and  $g_{i,t}^*$  be a  $k_i^* \times 1$  vector that contains country-specific foreign variables and the oil price. Following Pesaran et al. (2004), the country-specific foreign variables are constructed as a weighted average of other countries' corresponding domestic variables:

$$g_{i,j,t}^* = \sum_{l=0}^N w_{i,l} g_{l,j,t}, \quad i = 0, 1, 2, \dots, N,$$

where  $g_{i,j,t}$  ( $g_{i,j,t}^*$ ) is the element of  $g_{i,t}$  ( $g_{i,t}^*$ ) corresponding to variable  $j$ . Following the mainstream GVAR literature, we use a weighting scheme based on bilateral trade exposure (average trade shares over the period 2002 - 2008) because trade is a major driver of international transmission.<sup>1</sup> In particular,  $w_{i,l}$  is the ratio of

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<sup>1</sup>A matrix of trade weights used for constructing the country-specific foreign variables is available upon request.

the total trade between country  $i$  and country  $j$  to the total trade of country  $i$  with all of its trade partners, such that  $w_{i,i} = 0$  and  $\sum_{l=0}^N w_{i,l} = 1$ .

Our model uses the real effective exchange rate  $q_{i,t}$ , defined following Dees et al. (2007a) as

$$\begin{aligned} q_{i,t} &= \sum_{l=0}^N w_{i,l}(e_{i,t} - e_{l,t}) + p_{i,t}^* - p_{i,t} \\ &= \overleftarrow{e}_{i,t} - \overleftarrow{e}_{i,t}^*, \end{aligned}$$

where  $\overleftarrow{e}_{i,t} = e_{i,t} - p_{i,t}$ ,  $\overleftarrow{e}_{i,t}^* = e_{i,t}^* - p_{i,t}^*$ ,  $e_{i,t}$  is the nominal exchange rate in terms of domestic currency relative to the US dollar and  $p_{i,t}$  is the consumer price index (CPI).

Given the nature of macroeconomic variables, we treat all domestic variables as endogenous. For  $i = 1, 2, \dots, N$ , the set of endogenous variables is

$$g_{i,t} = (q_{i,t}, \pi_{i,t}, x_{i,t}, y_{i,t}, gov_{i,t}, open_{i,t}, si_{i,t})',$$

and the set of weakly exogenous variables is

$$g_{i,t}^* = (\pi_{i,t}^*, x_{i,t}^*, y_{i,t}^*, gov_{i,t}^*, si_{i,t}^*, oil_t)'$$

For  $i = 0$  (the US model), the set of endogenous variables is

$$g_{0,t} = (\pi_{0,t}, x_{0,t}, y_{0,t}, gov_{0,t}, open_{0,t}, si_{0,t}, oil_t)'$$

and the set of weakly exogenous variables is

$$g_{0,t}^* = (\overleftarrow{e}_{0,t}^*, \pi_{0,t}^*, x_{0,t}^*, y_{0,t}^*, gov_{0,t}^*)'$$

For all country-specific models except the US model, we include the real effective exchange rate ( $q_{i,t}$ ), the inflation rate ( $\pi_{i,t}$ ), the traded-nontraded productivity differential ( $x_{i,t}$ ), real output ( $y_{i,t}$ ), the government consumption share ( $gov_{i,t}$ ), the openness of the economy ( $open_{i,t}$ ) and the nominal interest rate ( $si_{i,t}$ ) as endogenous variables. All non-US models contain the foreign inflation rate ( $\pi_{i,t}^*$ ), the foreign traded-nontraded productivity differential ( $x_{i,t}^*$ ), the foreign real output ( $y_{i,t}^*$ ), the foreign government consumption share ( $gov_{i,t}^*$ ), the foreign nominal interest rate ( $si_{i,t}^*$ ) and oil prices ( $oil_t$ ) as weakly exogenous variables. Foreign government consumption is included as a proxy for foreign debt accumulation. We do not include  $open_{i,t}^*$  and  $q_{i,t}^*$  in  $g_{i,t}^*$  because they are highly correlated with  $open_{i,t}$  and  $q_{i,t}$ , which are already included in the set of endogenous variables. In the case of the US model, we do not impose the same specification as that for other countries. Specifically, the US foreign interest rate variable ( $si_{0,t}^*$ ) is not included in the US model because of the dominance of the US financial market in the global economy. The oil price variable ( $oil_t$ ) is included as endogenous in order to allow other variables to influence oil prices. As the US is the base economy,  $\overleftarrow{e}_{0,t}^*$  is treated as exogenous in its model.

## 4.1 Individual Country/Region-Specific Models

Unlike the existing GVAR literature, we estimate individual models on a country-group basis, instead of a country-by-country basis. This is because some of our country series are very short and panel data methods can improve the efficiency of our estimated coefficients. In particular, we divide our country set into four panels according to region and the level of economic development. It includes European, non-European developed, Asian developing and non-Asian developing country groups (with the exception of the US). See Table B1 for details of the countries in each group. Then we apply the within-group estimation technique to estimate each country group model separately.<sup>2</sup> This technique relies on the restrictions that the long-run and short-run coefficients in the model for each country from the same country group are the same, but it allows country-specific intercepts to be different across different countries. This assumption seems appropriate in light of the fact that each group consists of country members with similar features. Once we have estimated the model for each country group, we treat each country equation in that group as if it had been estimated separately. We estimate the US country-specific model separately because the US model has a different specification from that of other countries.

When constructing individual country models, we assume that all economies, with the exception of the US, are small relative to the world economy so that the foreign and global variables in each individual country model satisfy the assumption of weak exogeneity. This is the main requirement for estimation, when using the GVAR modelling approach. The validity of such an assumption depends on the relative sizes of the countries/regions in the global model and on whether the idiosyncratic shocks of the individual country models are cross-sectionally weakly correlated. Later, we discuss formal tests of weak exogeneity for foreign and global variables and provide empirical evidence on the test results.

We use augmented vector autoregressive (VARX\*) models with a second-order dynamic specification for domestic variables and a first-order dynamic specification for foreign variables in all individual country models. This choice was dictated by the small number of available time series observations and it corresponds to the GVAR literature, most of which allows a maximum lag of two in the model. The VARX\*(2,1) specification for each country  $i$  can be written as

$$g_{i,t} = h_{i,0} + \Phi_{i,1}g_{i,t-1} + \Phi_{i,2}g_{i,t-2} + \Psi_{i,0}g_{i,t}^* + \Psi_{i,1}g_{i,t-1}^* + u_{i,t} \quad (1)$$

for  $i = 0, 1, 2, \dots, N$ , and  $t = 1, 2, \dots, T$ , where  $g_{i,t}$  is a  $k_i \times 1$  vector of  $I(1)$  endogenous variables,  $g_{i,t}^*$  is

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<sup>2</sup>Alvarez and Arellano (2003) and Judson and Owen (1999) show that within-group estimation seems to have better performance than other methods for sizes of T and N close to our data set.

a  $k_i^* \times 1$  vector of  $I(1)$  weakly exogenous variables,  $h_{i,0}$  is a  $k_i \times 1$  vector of constant terms,  $\Phi_{i,1}$  and  $\Phi_{i,2}$  are  $k_i \times k_i$  matrices of coefficients associated with lagged endogenous variables,  $\Psi_{i,0}$  and  $\Psi_{i,1}$  are  $k_i \times k_i^*$  matrices of coefficients associated with weakly exogenous variables, and  $u_{i,t}$  is a  $k_i \times 1$  vector of idiosyncratic country-specific shocks. We assume that  $u_{i,t}$  is serially uncorrelated with mean zero and a nonsingular variance-covariance matrix  $\Sigma_{u,i}$ , i.e.  $u_{i,t} \sim iid(0, \Sigma_{u,i})$ . However, we allow these shocks to be contemporaneously correlated across countries or regions to a limited degree.

The corresponding conditional vector error correction model (VECMX\*) is given by

$$\Delta g_{i,t} = h_{i,0} - \Pi_i z_{i,t-1} - \Phi_{i,2} \Delta g_{i,t-1} + \Psi_{i,0} \Delta g_{i,t}^* + u_{i,t}, \quad (2)$$

where

$$\Pi_i = (I - \Phi_{i,1} - \Phi_{i,2}, -\Psi_{i,0} - \Psi_{i,1}) \text{ and } z_{i,t-1} = \left( g'_{i,t-1}, g_{i,t-1}^* \right)'$$

The cointegrating relationship among variables is summarized in a  $k_i \times (k_i + k_i^*)$  matrix  $\Pi_i$ . Suppose that the rank of  $\Pi_i$  is  $r_i \leq k_i$ , implying that there exist  $r_i$  long-run relationships among the domestic and country-specific foreign variables. The matrix  $\Pi_i = \alpha_i \beta_i'$ , where  $\alpha_i$  is a  $k_i \times r_i$  loading matrix of full column rank and  $\beta_i$  is a  $(k_i + k_i^*) \times r_i$  matrix of cointegrating vectors of rank  $r_i$ . In this paper, we use panel dynamic OLS (DOLS) to estimate cointegrating restrictions  $\widehat{\beta}_i$ , where the panel includes all countries in the same country group as country  $i$ . Conditional on  $\widehat{\beta}_i$ , we estimate the remaining parameters of the VECMX\* by using OLS.

Equation (1) can be rewritten in terms of  $z_{i,t}$  as

$$L_{i,0} z_{i,t} = h_{i,0} + L_{i,1} z_{i,t-1} + L_{i,2} z_{i,t-2} + u_{i,t}, \quad (3)$$

where

$$L_{i,0} = (I_{k_i}, -\Psi_{i,0}), \quad L_{i,1} = (\Phi_{i,1}, \Psi_{i,1}) \text{ and } L_{i,2} = (\Phi_{i,2}, \Psi_{i,2})$$

so that estimates of (3) can be deduced from estimates of (2). Note that  $\Psi_{i,2} = 0_{k_i \times k_i^*}$ , and  $L_{i,0}$ ,  $L_{i,1}$  and  $L_{i,2}$  are  $k_i \times (k_i + k_i^*)$  matrices with the condition that  $L_{i,0}$  has full row rank for  $i = 0, 1, \dots, N$ . Also note that any cross equation restrictions that are imposed on the panel version of (1) during estimation will imply corresponding structure on the  $L_{i,0}$ ,  $L_{i,1}$  and  $L_{i,2}$  for all  $i$  in the same country group.

## 4.2 Constructing the GVAR Model

Special techniques are needed to combine the individual country models into the GVAR model owing to the differences between the US model and the models for the remaining countries. Let  $k = \sum_{i=0}^N k_i$  be the total

number of endogenous variables in the global model and define a  $(k + 1) \times 1$  vector of the global variables,  $\overset{\circ}{g}_t$  as

$$\overset{\circ}{g}_t = (\overleftarrow{g}'_{0,t}, \tilde{g}'_{1,t}, \dots, \tilde{g}'_{N,t})',$$

with

$$\overleftarrow{e}_{0,t} = (\overleftarrow{e}_{0,t}, \pi_{0,t}, x_{0,t}, y_{0,t}, gov_{0,t}, open_{0,t}, si_{0,t}, oil_t)',$$

and

$$\tilde{g}_{i,t} = (\overleftarrow{e}_{i,t}, \pi_{i,t}, x_{i,t}, y_{i,t}, gov_{i,t}, open_{i,t}, si_{i,t})', \quad i = 1, 2, \dots, N.$$

As set out above, the total number of elements in  $\overset{\circ}{g}_t$  is one more than the total number of endogenous variables in the global model because  $\overset{\circ}{g}_t$  includes  $\overleftarrow{e}_{0,t}$  which is not an endogenous variable in the US model. Note that  $e_{0,t} = 0$  and hence  $\overleftarrow{e}_{0,t} = e_{0,t} - p_{0,t} = -p_{0,t}$ , which will be the first element of the vector  $\overset{\circ}{g}_t$ . It is easy to see that the vector  $z_{i,t}$  can be written in terms of  $\overset{\circ}{g}_t$  as

$$z_{i,t} = W_i \overset{\circ}{g}_t, \quad (4)$$

where  $W_i$  for  $i = 0, 1, 2, \dots, N$  are  $(k_i + k_i^*) \times (k + 1)$  matrices of fixed known constants called the "link" matrices, defined in terms of the country-specific weights. The link matrices play an important role in linking up all country-specific models to construct the GVAR model.

Substituting  $z_{i,t}$  from (4) into (3), we can write equation (3) as

$$L_{i,0} W_i \overset{\circ}{g}_t = h_{i,0} + L_{i,1} W_i \overset{\circ}{g}_{t-1} + L_{i,2} W_i \overset{\circ}{g}_{t-2} + u_{i,t}, \quad i = 0, 1, 2, \dots, N, \quad (5)$$

where  $L_{i,j} W_i$  are  $k_i \times (k + 1)$  matrices for  $j = 0, 1, 2$ . Stacking these equations yields

$$\overset{\circ}{H}_0 \overset{\circ}{g}_t = h_0 + \overset{\circ}{H}_1 \overset{\circ}{g}_{t-1} + \overset{\circ}{H}_2 \overset{\circ}{g}_{t-2} + u_t, \quad (6)$$

where

$$\overset{\circ}{H}_j = \begin{pmatrix} L_{0,j} W_0 \\ L_{1,j} W_1 \\ \vdots \\ L_{N,j} W_N \end{pmatrix} \text{ for } j = 0, 1, 2, \quad h_0 = \begin{pmatrix} h_{0,0} \\ h_{1,0} \\ \vdots \\ h_{N,0} \end{pmatrix}, \quad \text{and } u_t = \begin{pmatrix} u_{0,t} \\ u_{1,t} \\ \vdots \\ u_{N,t} \end{pmatrix}.$$

$\overset{\circ}{H}_j$  is a  $k \times (k + 1)$  matrix,  $h_0$  is a  $k \times 1$  vector of country-specific intercepts and  $u_t$  is a  $k \times 1$  vector of reduced-form residuals. To solve for the endogenous variables of the global model, we define the  $k \times 1$  vector  $g_t$  of global endogenous variables as

$$g_t = (\tilde{g}'_{0,t}, \tilde{g}'_{1,t}, \dots, \tilde{g}'_{N,t})',$$

where

$$\tilde{g}_{0,t} = (p_{0,t}, x_{0,t}, y_{0,t}, gov_{0,t}, open_{0,t}, si_{0,t}, oil_t)',$$

and

$$\tilde{g}_{i,t} = (\overleftarrow{e}_{i,t}, \pi_{i,t}, x_{i,t}, y_{i,t}, gov_{i,t}, open_{i,t}, si_{i,t})', \quad i = 1, 2, \dots, N.$$

The vector  $g_t$  differs from  $\mathring{g}_t$  because it does not include  $\overleftarrow{e}_{0,t}$ . Note that the first element of  $\mathring{g}_t$  is  $\overleftarrow{e}_{0,t}$  is equal to  $-p_{0,t}$ , so solving the GVAR model involves solving for the US price level  $p_{0,t}$  rather than inflation  $\pi_{0,t}$  in the vector  $g_t$  of global endogenous variables. We can write  $\mathring{g}_t$  in terms of  $g_t$  as

$$\mathring{g}_t = S_0 g_t - S_1 g_{t-1}, \quad (7)$$

where  $(k+1) \times k$  matrices of  $S_j$  for  $j = 0, 1$  are given by

$$S_0 = \begin{pmatrix} -1 & \mathbf{0}_{1 \times (k-1)} \\ 1 & \mathbf{0}_{1 \times (k-1)} \\ \mathbf{0}_{(k-1) \times 1} & I_{(k-1)} \end{pmatrix} \text{ and } S_1 = \begin{pmatrix} 0 & \mathbf{0}_{1 \times (k-1)} \\ 1 & \mathbf{0}_{1 \times (k-1)} \\ \mathbf{0}_{(k-1) \times 1} & \mathbf{0}_{(k-1) \times (k-1)} \end{pmatrix}.$$

Substituting  $\mathring{g}_t$  from equation (7) into (6), we have

$$H_0 g_t = h_0 + H_1 g_{t-1} + H_2 g_{t-2} + H_3 g_{t-3} + u_t, \quad (8)$$

where

$$H_0 = \mathring{H}_0 S_0, \quad H_1 = \mathring{H}_1 S_0 + \mathring{H}_0 S_1, \quad H_2 = \mathring{H}_2 S_0 - \mathring{H}_1 S_1, \quad H_3 = -\mathring{H}_2 S_1.$$

Then the GVAR model can be written as

$$g_t = a_0 + F_1 g_{t-1} + F_2 g_{t-2} + F_3 g_{t-3} + \varepsilon_t, \quad (9)$$

where  $F_j = H_0^{-1} H_j$  for  $j = 1, 2, 3$ ,  $a_0 = H_0^{-1} h_0$ , and  $\varepsilon_t = H_0^{-1} u_t$ .

## 5 Shock Identification Approach

### 5.1 Persistence Profiles

The persistence profiles developed by Pesaran and Shin (1996) use a variance-based approach for measuring the speed of convergence to long-run equilibrium after system shocks on the cointegrating relationships,

and hence they provide unique time profiles regardless of the prior orthogonalisation of the shocks. The persistence profiles also allow us to reassess the choice of the long-run relationships, and to check if an overidentifying restriction corresponding to a long-run relationship is actually valid. The intuition behind these checks is that if the choice of the long-run relationship is reasonable and its restriction is indeed a cointegrating vector, the persistence profile should approach zero at a reasonable speed as the horizon increases ( $h \rightarrow \infty$ ).

We assume that the GVAR model (9) has a fundamental moving average representation:

$$g_t = d_t + \sum_{j=0}^{\infty} B_j \varepsilon_{t-j}, \quad (10)$$

where  $d_t$  is the deterministic component of  $g_t$ . To obtain a  $k \times k$  matrix  $B_j$ , we apply Lemma 1 of Pesaran and Shin (1996) and compute  $B_j$  recursively as

$$B_j = F_1 B_{j-1} + F_2 B_{j-2} + F_3 B_{j-3}, \quad j = 1, 2, \dots, \quad (11)$$

where  $B_0 = I_k$  and  $B_j = 0$  for  $j < 0$ .

As before, we assume that there exist  $r_i$  long-run relationships among variables such that  $\beta'_i z_{i,t}$  is stationary. However, the set of variables in the GVAR model (9) is given by the vector  $g_t$  rather than  $z_{i,t}$ . Therefore we construct the cointegrating relationships,  $\beta'_i z_{i,t}$  using the link between these two vectors. We use equations (4) and (7) to link the vector  $z_{i,t}$  to the vector  $g_t$  and obtain

$$z_{i,t} = W_i \overset{\circ}{g}_t = W_i (S_0 g_t - S_1 g_{t-1}). \quad (12)$$

Then we can substitute (10) into (12) to obtain

$$z_{i,t} = W_i (S_0 d_t - S_1 d_{t-1}) + W_i S_0 B_0 \varepsilon_t + \sum_{j=1}^{\infty} W_i (S_0 B_j - S_1 B_{j-1}) \varepsilon_{t-j}. \quad (13)$$

Let  $\beta'_{j,i}$  be the  $j^{\text{th}}$  cointegrating relationship in country  $i$ . Pesaran and Shin (1996) define the persistence profile of  $\beta'_{j,i} z_{i,t}$  at horizon  $h$  (denoted by  $H(\beta'_{j,i} z_{i,t}; h)$ ) to be the difference between the conditional variances of the  $h$ -step-ahead and the  $(h-1)$ -step-ahead forecasts of a cointegrating relationship  $\beta'_{j,i} z_{i,t}$ , given the information set up to time  $(t-1)$ . The scaled measure of the persistence profile of  $\beta'_{j,i} z_{i,t}$ , denoted by a  $1 \times 1$  vector,  $PP(\beta'_{j,i} z_{i,t}; h)$  is

$$PP(\beta'_{j,i} z_{i,t}; h) = \frac{H(\beta'_{j,i} z_{i,t}; h)}{H(\beta'_{j,i} z_{i,t}; 0)}.$$

The value of the persistence profile is one on impact by construction. We can then use (13) to write the persistence profile of  $\beta'_{j,i}z_{i,t}$  with respect to a system shock to  $\varepsilon_t$  at horizon  $h$  as

$$PP(\beta'_{j,i}z_{i,t}; h) = \frac{\beta'_{j,i}W_iC_h\Sigma_\varepsilon C'_hW'_i\beta_{j,i}}{\beta'_{j,i}W_iC_0\Sigma_\varepsilon C'_0W'_i\beta_{j,i}}, \quad h = 0, 1, 2, \dots, \quad (14)$$

where  $\Sigma_\varepsilon$  is the variance-covariance matrix of  $\varepsilon_t$ , and

$$C_0 = S_0B_0 \text{ and } C_h = S_0B_h - S_1B_{h-1}.$$

## 5.2 Impulse Responses Using Sign Restrictions

Most existing GVAR literature uses generalized impulse response functions (GIRF). The GIRF approach considers shocks to an individual element in one country and integrates out the effects of other shocks using the observed error distributions, so it does not require any restrictions for the orthogonalisation of shocks. However, with non-orthogonal shocks, it is more difficult to interpret and understand the reasons behind the responses of each variable, which is a serious shortcoming when our interest is in the responses of variables to a particular underlying structural shock in a particular country.

We implement sign restrictions in the context of our global framework to identify simultaneous productivity improvement shocks in eight Asian emerging economies. With this approach, we can identify a particular economic shock by imposing a small set of a priori restrictions that are widely agreed upon by many economists. There are three main advantages of using this approach in the context of our global model. First, the sign restriction approach allows us to concentrate on identifying just a subset of structural innovations of interest. This is useful because the identification of all the shocks required for conventional identification schemes may be impossible in a multi-country model that includes a large number of endogenous variables. Second, one can easily increase the number of sign restrictions to help identify the structural shocks in question, when working with the global context. This leads to stronger results than those in the mainstream sign-restriction literature, most of which is based on VARs in a closed-economy setting. Third, we can obtain results that are robust to the ordering of the variables and the countries in the GVAR.

Consider the reduced-form VAR model (9) where  $\varepsilon_t = H_0^{-1}u_t$ , and  $u_t$  is a  $k \times 1$  vector of reduced-form errors. Our interest is to obtain impulse responses to economically meaningful structural shocks. We adopt a common assumption in the VAR literature that there are a total of  $k$  orthogonal structural innovations  $v_t$ , having unit variance. The relationship between reduced-form errors and structural shocks is given by  $v_t = A^{-1}u_t$  with  $\Sigma_u = E[u_tu'_t]$  and  $\Sigma_v = AA'$ . It is obvious that the identification of a  $k \times k$  matrix  $A$  is

necessary for achieving the independence of the fundamental innovations. We can use the moving average representation (10) to derive the  $k \times k$  matrix  $r(h)$  of impulse responses of all endogenous variables in the GVAR to  $k$  structural underlying shocks  $v_t$  at horizon  $h$  and obtain

$$r(h) = B_h H_0^{-1} A, \quad (15)$$

where the element in the  $s^{th}$  row and  $o^{th}$  column of the matrix  $r(h)$  represents the impulse response of the  $s^{th}$  endogenous variable to the structural shock to the  $o^{th}$  endogenous variable in the GVAR.

As mentioned earlier, the elements of  $v_{i,t}$  are uncorrelated but some elements of  $v_{i,t}$  and  $v_{j,t}$  with  $i \neq j$  may be correlated, or in other words, the shocks are orthogonal to other shocks within a given country but may be correlated with the shocks from different countries. However, the GVAR framework includes country-specific foreign variables in each individual country model to capture the cross-country correlations, and they are viewed as proxies for the common global factors. This leaves cross-sectionally weakly correlated remaining shocks across countries, and hence the interpretation of the shocks as country-specific seems to be reasonable. Later, we provide supporting empirical evidence on this issue.

We impose sign restrictions by using the procedures suggested by Rubio-Ramirez et al. (2010) and implemented by Eickmeier and Ng (2011) in the GVAR context. We compute the lower triangular Cholesky matrix  $\vec{A}_i$  of the covariance matrix of residuals for each country model  $i$  such that  $\Sigma_{u,i} = \vec{A}_i \vec{A}_i'$  where  $i = 0, 1, 2, \dots, N$  and  $i = 0$  for the US. Then we construct the  $k \times k$  matrix  $\vec{A}_{GV}$  as a block diagonal matrix with  $\vec{A}_i$  matrices on the diagonal blocks. This can be represented as

$$\vec{A}_{GV} = \begin{bmatrix} \vec{A}_0 & 0 & \cdots & 0 & 0 \\ 0 & \ddots & & & 0 \\ \vdots & & \vec{A}_i & & \vdots \\ 0 & & & \ddots & 0 \\ 0 & 0 & \cdots & 0 & \vec{A}_N \end{bmatrix}.$$

We note that the decomposition of the covariance matrix  $\Sigma_{u,i}$  to generate the matrix  $\vec{A}_i$  is not unique. In particular, we can find an infinite number of  $k_i \times k_i$  "candidate" matrices  $\tilde{A}_i$  such that  $\Sigma_{u,i} = (\vec{A}_i M_{k_i})(M_{k_i}' \vec{A}_i') = \tilde{A}_i \tilde{A}_i'$ , where  $\tilde{A}_i = \vec{A}_i M_{k_i}$ , and  $M_{k_i} = [m^{(1)}, m^{(2)}, \dots, m^{(k_i)}]$  is any  $k_i \times k_i$  orthonormal matrix. This underlies the implementation of sign restrictions. Suppose that a shock originates in country  $j$ . We randomly draw a set of  $k_j \times k_j$  orthonormal rotation matrices  $M_{k_j}$  and construct a "candidate" matrix  $\tilde{A}_{GV}$  as

$$\tilde{A}_{GV} = \begin{bmatrix} \vec{A}_0 & 0 & \dots & 0 & 0 \\ 0 & \ddots & & & 0 \\ \vdots & & \vec{A}_j M_{k_j} & & \vdots \\ 0 & & & \ddots & 0 \\ 0 & 0 & \dots & 0 & \vec{A}_N \end{bmatrix}.$$

The  $k \times 1$  vector of impulse responses of all endogenous variables in the GVAR to a shock to the  $a^{th}$  element of structural shock  $v_{j,t}$  corresponding to country  $j$  is given by  $r_{j_a}(h) = B_h H_0^{-1} \tilde{A}_{GV} c_{j_a}$  where  $c_{j_a}$  is a  $k \times 1$  selection vector with unity at the  $(\sum_{i=0}^{j-1} k_i + a)^{th}$  element and zeros elsewhere. We examine a wide range of possible choices for  $\tilde{A}_{GV}$  by repeatedly drawing sets of orthonormal rotation matrices  $M_{k_j}$  and checking corresponding impulse responses for each draw. If the candidate impulse responses satisfy a set of a priori identifying restrictions on the impulse response functions, we retain  $M_{k_j}$ . Otherwise, we discard  $M_{k_j}$ . We repeat this procedure until we retain 200 valid  $M_{k_j}$ s. Then we find the 16 percent, 50 percent and 84 percent quantiles of the distribution for the points on the impulse response functions.

Our method of computing the effects of simultaneous productivity shocks to eight countries is similar to that used when computing the effect of a single shock. We define a composite shock as a linear combination of variable-specific shocks across particular countries, and compute the responses of the composite shock as weighted averages of the responses to variable-specific shocks in given countries, using PPP-adjusted GDP weights that reflect the relative importance of each country within the world economy. Since we are interested in the effects of simultaneous productivity shocks to eight countries, we implement the sign restrictions by rotating eight matrices  $\tilde{A}_j$  (where  $j$  represents the eight countries that experience the shock) using a wide set of eight orthonormal matrices  $M_{k_j}$  to construct many possible choices of  $\tilde{A}_{GV}$ . Then we compute the impulse responses of the composite shock using  $r^{cp}(h) = B_h H_0^{-1} \tilde{A}_{GV} c^{cp}$  where  $c^{cp}$  is a  $k \times 1$  selection vector that picks out sign restricted responses to the eight shocks with suitable weights.

Moreover, we consider the forecast error variance decomposition indicating the contribution of the  $a^{th}$  element of the structural innovation  $v_{j,t}$  to the variance of the h-step ahead forecast error for the  $c^{th}$  element of  $g_t$ . The standard approach for the calculation of the forecast error variance decomposition is based on a set of orthogonalized shocks, and this is not appropriate for our GVAR model since the shocks across countries are not orthogonal. Therefore, we combine sign restrictions and a generalized forecast error variance decomposition to compute a scaled version of forecast error variance. More precisely, the proportion of the forecast error variance for the  $s^{th}$  element of  $g_t$  which can be explained by the  $a^{th}$  element of the structural innovation  $v_{j,t}$  at horizon  $h$ ,  $\phi_{j_a,s}(h)$  is given by

$$\phi_{j_a,s}(h) = \frac{(c'_s B_h H_0^{-1} \tilde{A}_{GV} c_{j_a})^2}{c'_s B_h H_0^{-1} \Sigma_u H_0^{-1'} B'_h c_s},$$

where  $c_s$  is a  $k \times 1$  selection vector in which the  $s^{th}$  element of  $g_t$  is one and all other elements are zeros and  $c_{j_a}$  is a  $k \times 1$  selection vector with unity at the  $(\sum_{i=0}^{j-1} k_i + a)^{th}$  element and zeros elsewhere.

### 5.2.1 Shock Identification

Our identification strategy simply requires restrictions on the short-run impulse responses of a small set of variables whose responses to a given shock are unambiguous according to theoretical predictions and conventional views. The responses of variables that can be subject to debate are left unconstrained.

The restrictions used for identifying a shock to productivity improvement in the traded sector are derived from the production function. The shock can be viewed as a supply shock. Thus, we consider productivity shocks in traded sectors as shocks that not only raise the traded good vs non traded good productivity differential on impact but which also raise real output on impact. Note that although an increase in the traded good vs non traded good productivity differential could possibly be induced by a decline in productivity in non-traded goods, one would expect real output decline in that case, and our restriction that output must increase precludes this scenario.

Our restrictions set the scene for examining the BS hypothesis, that predicts that productivity gains in the traded sector will lead to higher economic growth and an appreciation of the real exchange. Our restrictions allow both the traded good productivity and output to rise, but leave the real exchange unrestricted so that we can examine its response. There are several debates regarding real exchange rate behaviour in fast growing countries in Asia, and it is this that motivates us to examine the effects of simultaneous productivity shocks in China, India, Indonesia, Thailand, Philippines, Malaysia, Pakistan, and Sri Lanka, which are now emerging and developing economies in Asia. To model this composite shock, our short-run restrictions are imposed simultaneously on the impulse responses of productivity and real output for these eight countries.

## 6 Empirical Investigation

### 6.1 Preliminary Analysis

This paper extends the analysis in Dumrongrittikul (2011) that examined the effects of shocks in these four separate country groups (European, non-European developed, Asian developing and non-Asian developing country groups), into one that is undertaken in a GVAR setting, allowing for an understanding of the responses of variables in each country to shocks that can originate from anywhere in the world economy.

Using the results of unit root tests presented in this previous work, we treat all domestic and foreign variables - the real effective exchange rate ( $q_{i,t}$ ), the inflation rate ( $\pi_{i,t}$ ), the traded-nontraded productivity differential ( $x_{i,t}$ ), real output ( $y_{i,t}$ ), the government consumption share ( $gov_{i,t}$ ), the openness of the economy ( $open_{i,t}$ ), the nominal interest rate ( $si_{i,t}$ ) and oil prices ( $oil_t$ ) - as I(1) except for measures of domestic inflation in Asian developing countries and in the US that are considered as I(0).

Among this set of variables, Dumrongrittikul (2011) tested each possible theoretical long-run relationship including PPP, the Fisher equation, output convergence, UIP and a real exchange rate relationship derived from the BS hypothesis and the Edwards (1989) model separately by using Pedroni (1999) panel cointegration tests for each country group. The test results show that three long-run relationships - the Fisher equation, UIP and the real exchange rate relationship - hold. Then they considered two possible choices of over-identifying long-run restrictions that are (1) coefficients suggested by economic theory; and (2) cointegrating vectors estimated by panel DOLS, developed by Mark and Sul (2003). Note that with this method, the long-run restrictions for each country in the same country group are restricted to be the same.

In this paper, we follow Garratt et al. (2006) and Smith and Galesi (2011) i.e. the final choice of long-run relationships and their cointegrating vectors in the four country group models and in the US model used in our GVAR is based on a satisfactory performance of the GVAR model in terms of its stability, assured by examining the eigenvalues of the GVAR model, persistence profiles (convergence to zero at reasonable speed) and the shape of the impulse response functions. Based on these criteria, the final set of long-run relationships for each country group and the US model is summarized and presented in Table 1.

Table 1: Over-identifying long-run relations for each individual model

Country groups	Fisher	UIP	BSH and Edwards model	Inflation
European countries	-	$si_{i,t} - 0.641si_{i,t}^*$	$q_{i,t} + 0.439(x_{i,t}^* - x_{i,t}) - 0.145(y_{i,t}^* - y_{i,t}) + 0.137open_{i,t} - 0.683(gov_{i,t}^* - gov_{i,t})$	-
Non-European developed countries	-	$si_{i,t} - 0.940si_{i,t}^*$	$q_{i,t} + 0.247(x_{i,t}^* - x_{i,t}) + 0.110(y_{i,t}^* - y_{i,t}) - 0.685open_{i,t} + 0.213(gov_{i,t}^* - gov_{i,t})$	-
Asian developing countries	-	$si_{i,t} - si_{i,t}^*$	$q_{i,t} + 0.539(y_{i,t}^* - y_{i,t}) - 0.071open_{i,t} - 0.276(gov_{i,t}^* - gov_{i,t})$	$\pi_{i,t}$
Non-Asian developing countries	$si_{i,t} - \pi_{i,t}$	-	$q_{i,t} - 0.119(x_{i,t}^* - x_{i,t}) - 0.302(y_{i,t}^* - y_{i,t}) - 0.048open_{i,t} - 0.425(gov_{i,t}^* - gov_{i,t})$	-
US	-	-	-	$\pi_{0,t}$

Notes: An increase (a fall) in the value of the real exchange rate  $q_{i,t}$  represents real depreciation (appreciation).

The BSH and Edwards relationships reported in Table 1 suggest that an increase in the degree of openness in most country groups is associated with a real depreciation and higher government consumption is associated with a real appreciation after controlling for other factors. Note that we add the last column

named inflation because the data suggests that inflation in Asian developing countries and in the US is stationary and therefore can be viewed as being cointegrated with itself. We estimate individual models for each country group employing the set of over-identifying long-run restrictions, and then we separate them out into individual country-specific models prior to undertaking impulse response analysis.

## 6.2 Contemporaneous Effects of Foreign Variables on their Domestic Counterparts

Table B2 in Appendix B reports the estimates of impact elasticities of foreign variables on corresponding domestic variables, derived from the coefficient estimates associated with first differenced contemporaneous foreign variables,  $\Psi_{i,0}$  in the model (2). These elasticities measure the percentage change in a domestic variable caused by a one percent change in its corresponding foreign counterpart, and they show general co-movement among variables across different countries.

Our results show that all the elasticities relating to inflation and nominal interest rates are statistically significant, indicating strong linkages between financial variables across different regions. Looking at short-term interest rates, all contemporaneous elasticities have a positive sign. This implies co-movements between financial markets and cross-country monetary policy reactions, i.e. a change in the monetary policy employed by other countries may influence monetary policy in the domestic country within the same period. In particular, a 1 percent increase in foreign nominal interest rates can lead to a rise of 0.9 percent in the nominal interest rates in non-Asian developing countries.

The estimates of foreign inflation suggest that there is co-movement in inflation across countries. Impact elasticities are greater in developing countries than those in developed countries, revealing a stronger effect of foreign inflation on a change in domestic inflation in developing countries. However, the estimated elasticities of non-Asian developing countries have an unexpected sign, probably because these countries have very flexible monetary policy to prevent foreign inflation pass-through. For the European group and the US, we find low elasticities for inflation, suggesting that prices in the European group and in the US are not greatly affected by changes in foreign inflation in the short run. This finding is in line with the general idea that the transmission channel of inflation from large to small economies is unidirectional.

The estimates of foreign government consumption shares are positive and significant in all but one case, suggesting co-movement in fiscal policy across different countries. Most elasticities of productivity differentials and real GDP are small and insignificant. This is consistent with our expectation that real variables will react quite slowly to a change in foreign variables.

### 6.3 Weak Exogeneity of Foreign Variables

We test the validity of weak exogeneity assumption for the foreign variables and the oil price. Following Johansen (1992) and Granger and Lin (1995), this assumption means that there is no long-run force from  $g_{i,t}$  to  $g_{i,t}^*$ , regardless of lagged short-run effects between these variables. We can indirectly check this assumption by testing the joint significance of the estimated error correction terms in the auxiliary models for the country-specific foreign variables,  $g_{i,t}^*$  (See Dees et al. (2007b) for details). Note that we test the weak exogeneity on a country-group basis. Table B3 in Appendix B provides the test results. It shows that most of the foreign variables and the oil price can be considered as weakly exogenous in the individual models. This hypothesis is rejected only for inflation in the non-Asian developing country group, but this single result is unconvincing given the other results. In addition to our formal tests we also inspected the average pair-wise cross-section correlations for the endogenous variables and then the corresponding cross-section correlations in individual country VECMX\* residuals. We found substantial cross-section correlations between the endogenous variables, but the cross-section correlations of the associated residuals were much lower, especially in the case of output where the correlations fell from 92 percent to 2 percent on average. While the observed patterns for real exchange rates provided an exception (as was also the case in Dees et al (2007b)), these findings were consistent with a conclusion that the inclusion of foreign variables in the models helped to capture the cross-section correlations and common effects across countries, leaving the models with just a modest degree of correlation across the idiosyncratic shocks. Further details are available upon request.

### 6.4 Persistence Profiles

We compute the time profiles of the effects of system shocks on the cointegrating relationships that correspond to a GVAR model that imposes the long-run restrictions shown in Table 1. There are three types of long-run restrictions examined: the Fisher equation, UIP and the real exchange rate relationship.

The Fisher relation holds in the non-Asian developing country group only. As shown in Figure A1, it is obvious that the Fisher relations converge to equilibrium very quickly, within 2 years of a shock for all countries in this group. This corresponds to the results of Dees et al. (2007a). Compared to the Fisher relation, the UIP relations in Figure A2 are quite persistent, and they can deviate from equilibrium for a longer time after a shock. In particular, they require about 5 years for the European group and 7 - 8 years for the non-European developed and Asian developing country groups to return to equilibrium, which is similar to the findings of Lothian and Wu (2011).

Figure A3 shows that after a shock, the real exchange rate movements implied by our empirical model are quite persistent, especially in the Asian developing countries, where they last for well over ten years. This may be because several assumptions that underly the Balassa-Samuelson hypothesis are not perfectly satisfied; for example, imperfect competition, foreign exchange market intervention, trade barriers or high transaction costs are all features of Asian developing countries.

As mentioned earlier, the persistence profiles can be used to examine the choice of cointegrating vectors that identify the long-run relations. Since Figures A1 - A3 show that all persistence profiles are well behaved and eventually return to zero, it appears that our choice of the long-run relations in Table 1 is appropriate.

## 6.5 Impulse Response Analysis

We compute the impulse responses of each country in the GVAR using sign restrictions to identify a composite shock to productivity differential between traded and nontraded sectors in eight developing economies (China, India, Indonesia, Thailand, Philippines, Malaysia, Pakistan, and Sri Lanka) in Asia. Given space considerations, we show the responses for the main countries of interest and the aggregate responses for other countries, computed by weighting the impulse responses of individual countries using the averages of PPP-adjusted GDP over the period 2002 - 2008. Figures A4 - A11 display the median responses, together with two lines that show the 16<sup>th</sup> and 84<sup>th</sup> percentile impulse responses of the posterior distribution of the impulse response functions for each horizon.<sup>3</sup> In our analysis below, the word "significance" is used whenever both the 16<sup>th</sup> and 84<sup>th</sup> percentile responses are either above or below zero at a particular horizon.

We also note that our use of the median responses can be challenged, as Fry and Pagan (2011) point out that the median of the impulse responses will typically be generated by different models and there is nothing to ensure that the identified shocks are orthogonal. To examine this issue, we use Fry and Pagan's (2011) approach to find a single model which generates impulse responses that are as close to the median responses as possible, and compute impulse responses on the basis of the single model. The responses based on this approach, displayed by the dashed line in Figures A4 - A11, show that in general they are not statistically different from the median responses as they fall within the 16<sup>th</sup> and 84<sup>th</sup> percentiles of the posterior distribution.

### 6.5.1 Domestic Responses to Simultaneous Productivity Improvement Shocks

Figure A4 displays the impulse responses of domestic variables to a temporary composite shock to sectoral productivity differentials in developing Asia. This shock causes productivity in the traded sector relative to

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<sup>3</sup>Note that by construction, these error bands actually reflect model uncertainty, rather than sampling uncertainty.

the nontraded sector to initially rise in China (by 0.42 percent), in India (by 0.22 percent) and in Indonesia (by 0.15 percent) and other developing Asian countries by lower amounts. The shock is constrained to have an initial positive effect on real GDP in the eight developing countries in Asia, but the effects on real GDP in these countries are positive and statistically significant for longer, particularly in Pakistan and Sri Lanka, where effects are still strong after three years.

According to the BS model, productivity gains in the domestic traded sector relative to the nontraded sector raise real wages in both sectors due to domestic labour mobility. This induces an increase in the prices of domestic nontraded goods relative to domestic prices of traded goods, implying a real appreciation. We find that although the effect of a positive productivity shock in the traded sector is only short-lived, the real exchange rate responses in these eight countries seem to reflect the BS hypothesis. On the one hand, higher productivity improvement in traded sectors in eight Asian developing countries relative to the US and the other countries is followed by real appreciation of their currency against the US dollar in the medium term. Further, the responses of an overall domestic price are also in line with the transmission channel of the Balassa-Samuelson effect that a productivity gain in the traded sector will bring an increase in domestic prices due to higher prices of nontraded products and positive income effects. As shown in Figure A4, the shocks cause inflation in these countries to increase and this response is especially strong in China.

In addition, our findings suggest positive effects of the productivity shocks on trade related activities in these countries. This is what we expect as we normally think that technological progress leads to an efficient allocation of traded sector production at a global level. The countries that the shocks directly affect will have a competitive advantage as exporters and foreign-invested firms will also likely shift production to them. Exports from these countries will rise, as will imports of materials and components used in their production. Note that the responses of the ratio of international trade to GDP in China and India are small and insignificant, probably because these two countries are large and relatively less dependent on foreign countries.

The responses of short-term interest rates and government consumption shares in the eight countries are generally not significant, suggesting that monetary and fiscal policy makers do not respond to productivity shocks.

### **6.5.2 International Responses to Simultaneous Productivity Improvement Shocks**

Figures A5 - A11 show how traded sector productivity improvement shocks in Asian developing countries propagate internationally. Interestingly, we find that the traded-sector productivity shocks drive up productivity in the nontraded sector relative to that in the traded sector in almost all other countries including

the US, non-European developed countries and non-Asian developing countries as shown in Figure A6. This suggests that traded-sector productivity increases in Asia may make the firms that produce traded goods in the US and in other economies less internationally competitive as exporters. Many of these firms are in the manufacturing sector. They may move production facilities to Asian developing countries, and shift their productions and resources away from traded sectors toward nontraded sectors, inducing productivity improvement in their nontraded sectors. As a result, the traded good production in the countries that compete with exports in Asian developing countries can fall in the short run. Moreover, a rise of international trade shares in most countries may also be a result of an increase in their imports via the global supply chain.

In line with the BS hypothesis, US productivity growth in nontraded sectors after the shocks leads to a real depreciation of the US dollar against other currencies. In other words, Figure A5 shows that in the medium run all currencies appreciate against the US dollar in real terms. The shocks spill across borders and have different effects on the aggregate outputs of each country. That is, after the shocks real GDP declines in the US and non-European developed countries. On the other hand, they cause real GDP to increase in European and non-Asian developing countries but by less than that in Asian developing countries. See Figure A7.

Consider the consequences of the Asian developing country traded good productivity shocks on the US economy. After the shocks, US output in the traded sector falls and an associated real depreciation of the US dollar will decrease real wealth, raising prices (inflation) for US consumers, lowering economic welfare, and finally reducing US aggregate demand and output in the short run. The effect of foreign exchange markets on the volume of exports and imports (or on aggregate output) seems to be consistent with a J-Curve effect. In particular, a real depreciation of the US dollar after the shock will cause exports from the US to be relatively cheap and exports from Asian-developing countries to be relatively expensive. However, the volume/demand and the prices of imports and exports are normally slow to adjust. After some time, changes in real exchange rates can affect the prices of traded goods and services. As a result, the US production of traded sectors can compete with those in Asian developing countries and this may boost US exports to these countries. The US traded firms will produce more output and cause aggregate output to return slowly to the original level. See Figure A6. In addition, the responses of US short-term interest rates are in line with an economist's perspective. That is, when the depreciation of the US dollar improves the US trade balance/current account, this equates to capital outflows from the US to Asian developing countries in balance of payment accounts. Lower demand for US financial assets will put upward pressure on US interest rates.

In general, the traded good productivity shocks in Asia have little effect on government consumption

shares and short-term interest rates in other countries, suggesting no policy response to the shocks. An exception is found in non-Asian developing countries which experience a sudden drop of short-term interest rates in response to traded good productivity shocks in developing countries in Asia (see Figure A8).

## 6.6 Forecast Error Variance Decomposition Analysis

Given the identified productivity shocks, we determine the fractions of forecast error variance in real exchange rates are accounted for by these shocks. Table B4 shows, on the basis of the median estimates, forecast error variance shares of each country's real exchange rates explained by simultaneous productivity shocks to eight Asian developing countries at the selected year horizons,  $h = 0, 1, 2, 4,$  and  $8$ . We find that the forecast error variance shares explained by the shocks follow a similar path in most countries, although the shocks account for different proportions of their forecasted real exchange rate variations. See Figure A12. In particular, the productivity shocks explain a small proportion of forecast error variance shares of real exchange rates on impact. The explained shares rise strongly and reach a peak at the one-year or two-year horizon, suggesting that the shocks take some time to have their maximum impact on real exchange rate movements. The shocks make a smaller contribution to exchange rate fluctuations at longer horizons, reflecting their temporary effect on real exchange rates.

The forecast error variance shares explained are varied across different countries and horizons. The shares due to simultaneous productivity gains in Asian developing countries are quite limited in all countries in our sample, partly because their effects of the exchange rates associated in the countries in which shocks originated are small (eg only 1.9% of the movements in China's real exchange rate at the one-year horizon can directly be ascribed to the productivity shocks) and partly because these eight countries have small direct trade exposures to other economies and they are less internationally financially integrated, when compared to the US. As a result, the shocks originating in these countries are only weakly transmitted across borders, inducing only a small proportion of the forecast error variance of real exchange rates in foreign countries at any horizon.

## 6.7 Robustness Analysis

We perform two robustness checks to examine whether our results are sensitive to shock identification schemes. First, we consider the effects of simultaneous shocks to traded-sector productivity in China, India and Indonesia which are the three largest emerging countries in Asia. The shocks to these three countries constitute a subset of the shocks that we originally considered. We find that impulse responses are similar to the benchmark impulse responses to productivity shocks in eight Asian developing countries with respect

to sign, size and shapes, and their error bands become wider than the benchmark error bands. The results correspond to our expectation, as identifying shocks to the three countries involves a smaller set of sign restrictions, inducing a wider range of admissible responses. Second, we impose alternative and equally plausible identifying restrictions by imposing an additional restriction that productivity shocks will cause openness to increase in eight Asian developing countries. Our results are very robust to the new set of restrictions.

Generalized impulse response functions (GIRFs) are commonly used in the GVAR literature, and we compare these with our sign restricted responses and comment on some differences. Since the definitions of shocks differ in each case we do not expect GIRFs and sign-restricted impulses to be the same, but it is interesting to note that GIRFs suggest that positive shocks to the traded non-traded productivity differentials in Asian countries cause depreciation, whereas sign restricted impulse response suggest appreciation as predicted by economic theory. Further analysis of GIRFs implied by our GVAR delivers other unexpected findings,<sup>4</sup> and we attribute these to the non-orthogonal nature of shocks that are identified as GIRFs are calculated, and the corresponding difficulty of assigning any economic interpretation to such shocks.

## 7 Conclusion

This paper constructs a multi-country model covering sixteen developed and fourteen developing countries to investigate the international effects of traded sector productivity gains in eight Asian emerging countries (China, India, Indonesia, Thailand, Philippines, Malaysia, Pakistan and Sri Lanka) on key macroeconomic variables, particularly real exchange rates. We use an econometric strategy based on a global vector autoregression (GVAR), which captures direct channels of international shock transmission by including foreign variables in the models and indirect channels through error spillover effects, while keeping dimensionality manageable. The GVAR enables us to model international linkages between a large number of countries, which is an important feature given the multilateral nature of the real exchange rate.

We identify the structural shocks of interest by using sign restricted impulse response functions rather than generalized impulse response functions. This identification scheme is appropriate for our study because it gives our shocks a clear economic interpretation. We can also impose a large number of economic theory-based sign restrictions on the GVAR, yielding results that are robust to the ordering of variables and countries in the system.

Several results stand out. First, real exchange rate responses seem to reflect the Balassa-Samuelson

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<sup>4</sup>Readers can contact the authors for a set of diagrams that illustrate the empirical differences between GIRFs and sign restricted responses.

hypothesis, even though the productivity differentials between traded and nontraded sectors only change temporarily. In particular, traded-sector productivity gains in Asian developing economies are followed by medium-run appreciations of their domestic currencies. The real values of currencies in most non-US developed and other developing countries also appreciate in the short run. Second, the shocks drive up productivity in the nontraded sector relative to the traded sector in several other countries, including the US, non-European developed and non-Asian developing countries. This suggests that as more traded products are made in Asian developing countries, there is a compositional shift in the other countries' production and resources away from the traded sectors toward the nontraded sector, allowing their productivity in the nontraded sector to increase. Third, the shock to traded-sector productivity is a contributor to real GDP growth and leads to short-run increased inflation in Asian developing countries. However, the shocks spill over into foreign economies in different ways, inducing either positive or negative effects on the aggregate output of these other countries in the medium run. Fourth, traded sector productivity gains raise international trade in most developing Asian and developed countries as expected. Last, according to variance decomposition analysis, the role of the shocks in explaining real exchange rate fluctuations varies across countries. Indeed it appears that traded-sector productivity improvements in Asian emerging economies have a relatively small effect on real exchange rate fluctuations in both domestic and foreign countries at all horizons (relative to other shocks that influence exchange rates).

In reality, many types of shocks influence all economies around the world, and these are likely to contaminate our identification of productivity shocks to the traded sector and our understanding of their transmission. A more realistic application would involve a larger number of sign restrictions and identified shocks to capture more transmission channels than those modelled in this paper. It might be worth constructing more complex global models that include a larger set of possible real exchange rate determinants and a richer shock structure. This will be left for future research.

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# A Figure Appendix

Figure A1: Persistence profiles of system shocks on the Fisher relations

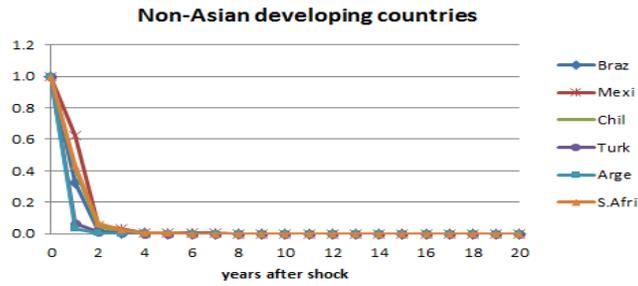


Figure A2: Persistence profiles of system shocks on the UIP relations

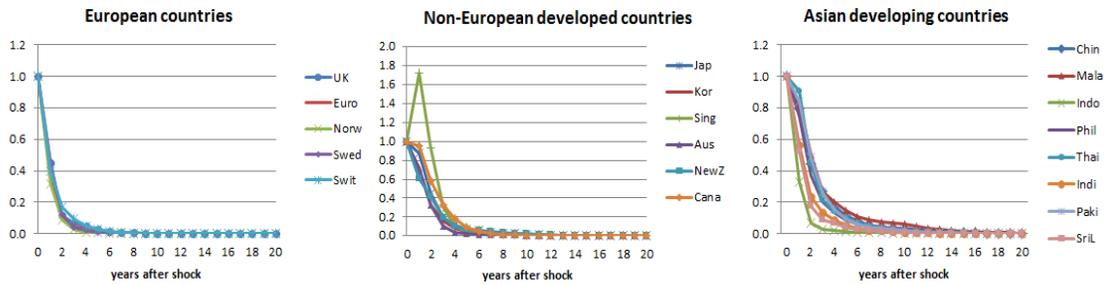


Figure A3: Persistence profiles of system shocks on the real exchange rate relations

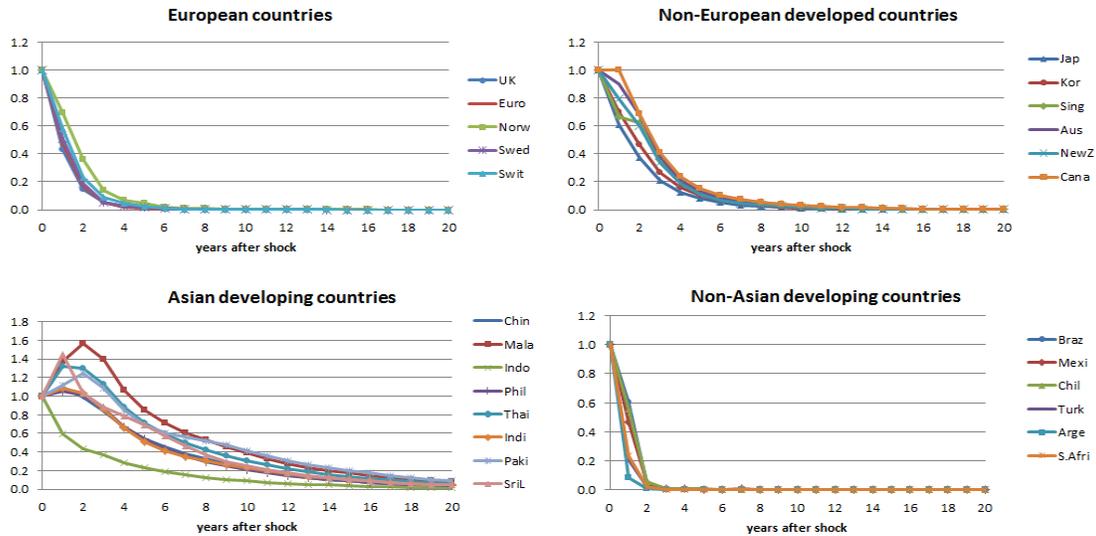
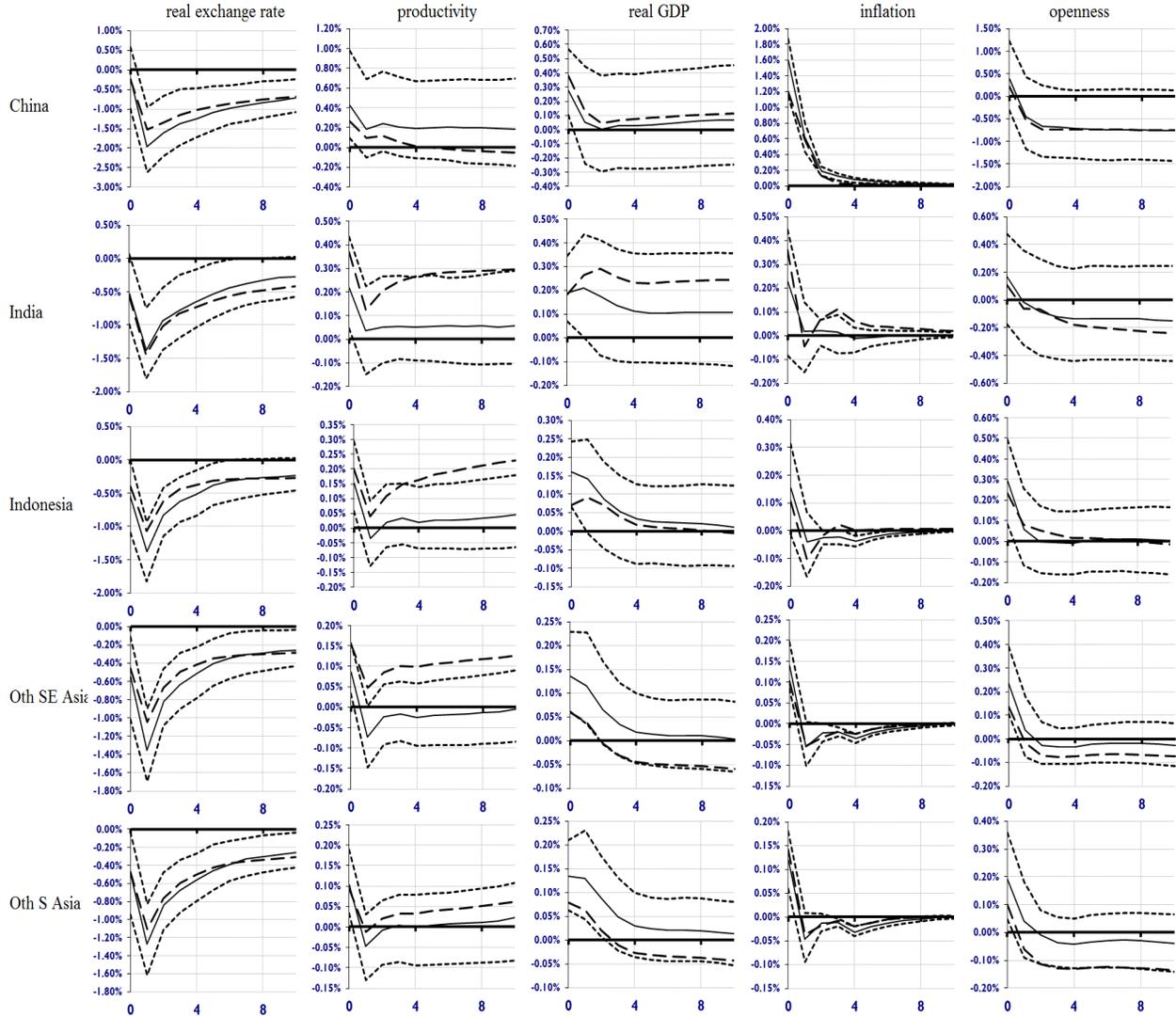
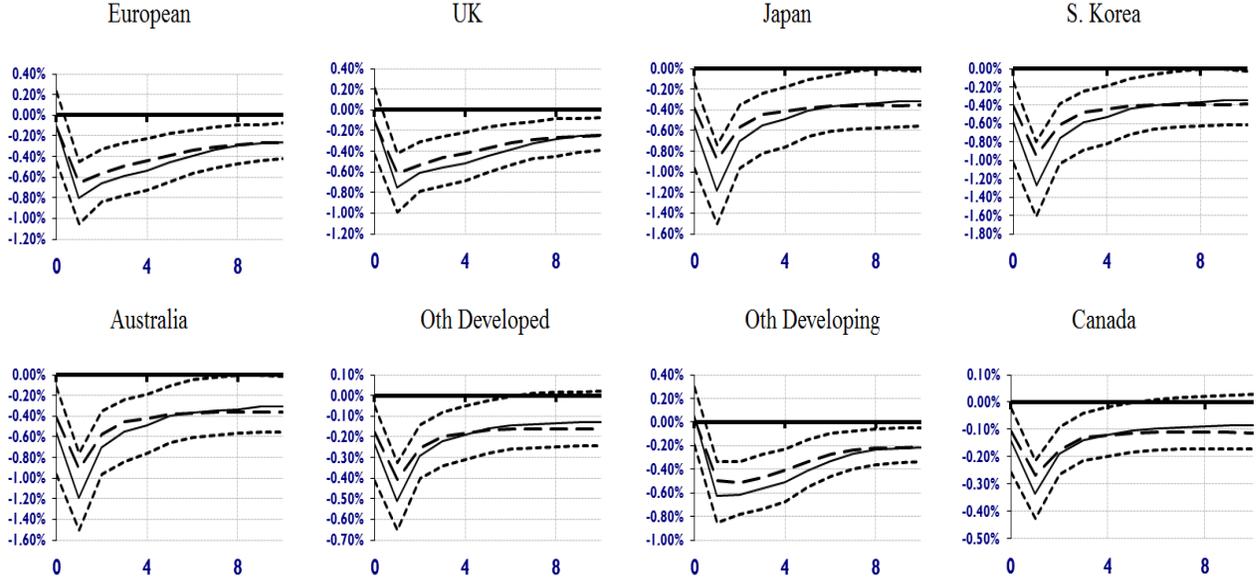


Figure A4: Transmission of a composite shock to traded-sector productivity improvement to selected variables in Asian developing countries



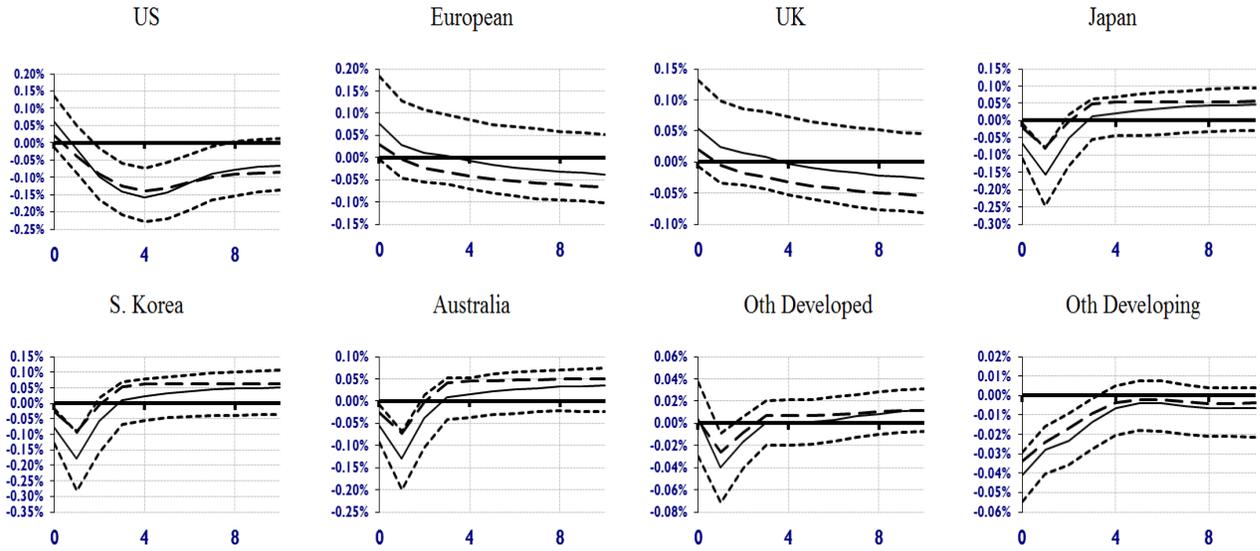
Notes: Solid line represents median impulse responses, and dashed line displays impulse responses implied by a single model as proposed by Fry and Pagan (2011). Dotted lines correspond to 16<sup>th</sup>/84<sup>th</sup> percentile responses. Note that 1) Other Southeast Asia countries include Thailand, Malaysia, and Philippines, 2) Other South Asia countries include Pakistan and Sri Lanka, 3) An increase (a fall) in the real exchange rate ( $q_t$ ) represents a real depreciation (appreciation).

Figure A5: Real exchange rate responses to a composite shock to traded-sector productivity improvement



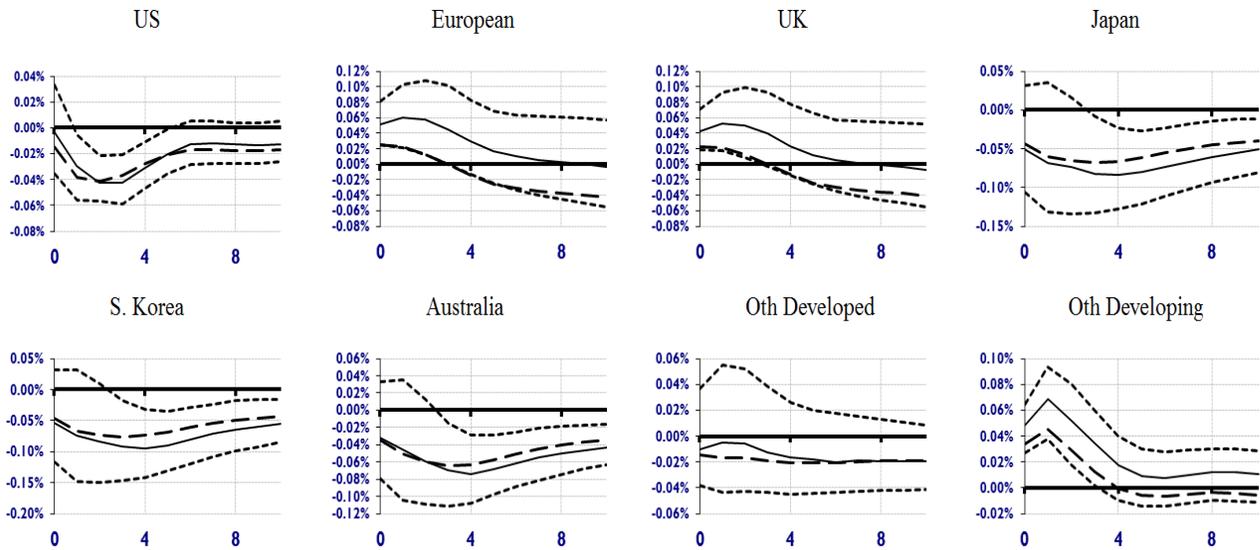
Notes: An increase (a fall) in the real exchange rate ( $q_t$ ) represents a real depreciation (appreciation). Solid line represents median impulse responses, and dashed line displays impulse responses implied by a single model as proposed by Fry and Pagan (2011). Dotted lines correspond to 16<sup>th</sup>/84<sup>th</sup> percentile responses. Note that 1) European countries include the Euro Area countries, Norway, Sweden, Switzerland, 2) Other developed countries include Canada, New Zealand, Singapore, 4) Other developing countries include Brazil, Mexico, Chile, Turkey, Argentina, South Africa.

Figure A6: Traded-nontraded productivity differential responses to a composite shock to traded-sector productivity improvement



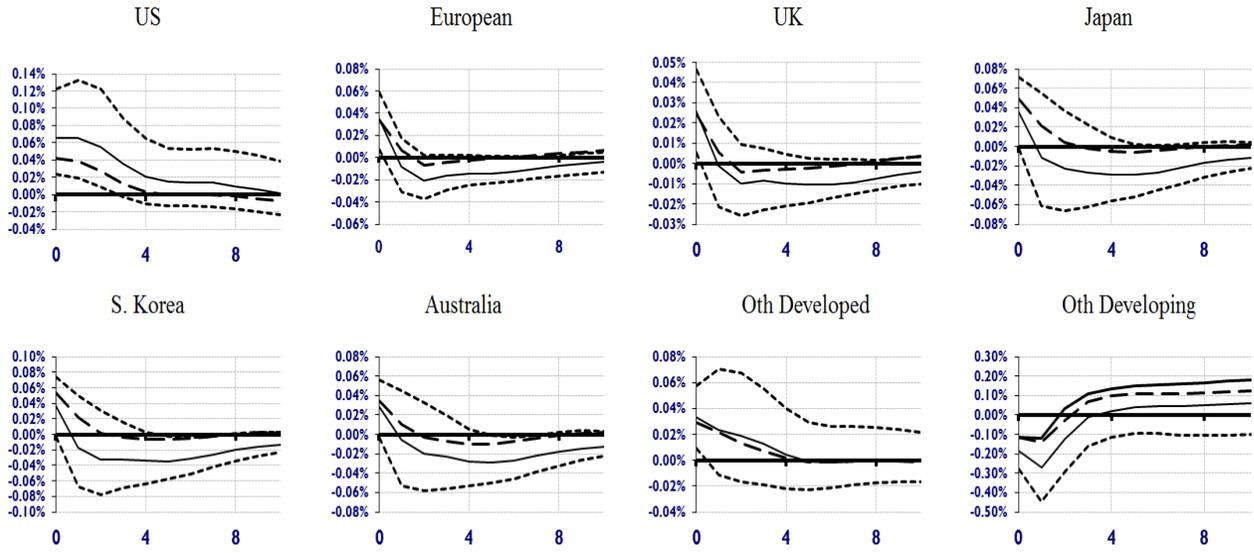
Notes: See Figure A5.

Figure A7: Real GDP responses to a composite shock to traded-sector productivity improvement



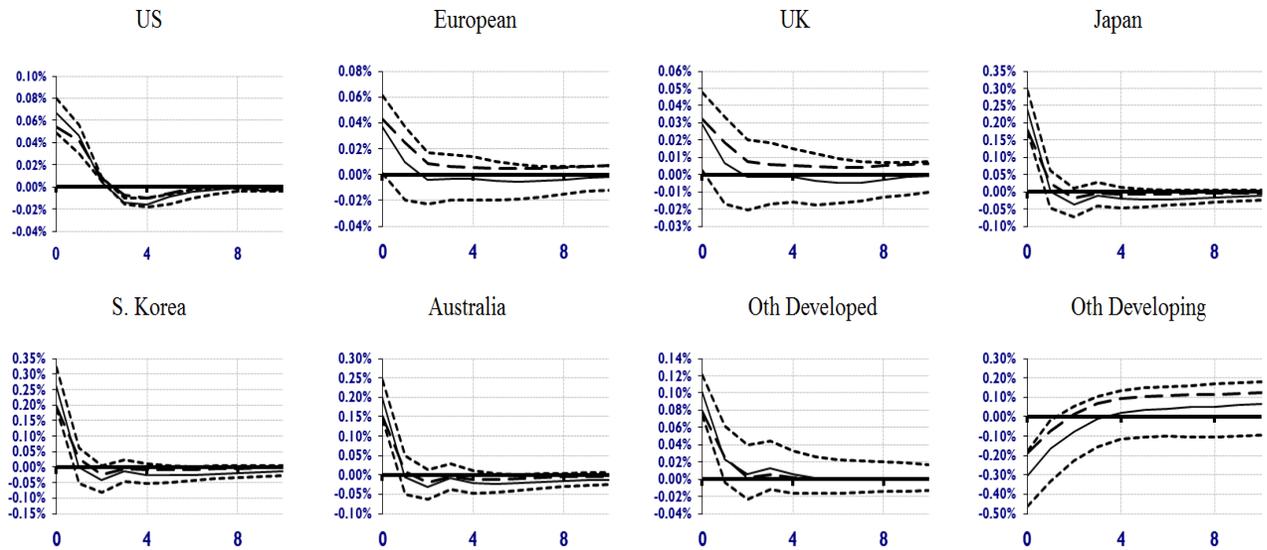
Notes: See Figure A5.

Figure A8: Short-term interest rate responses to a composite shock to traded-sector productivity improvement



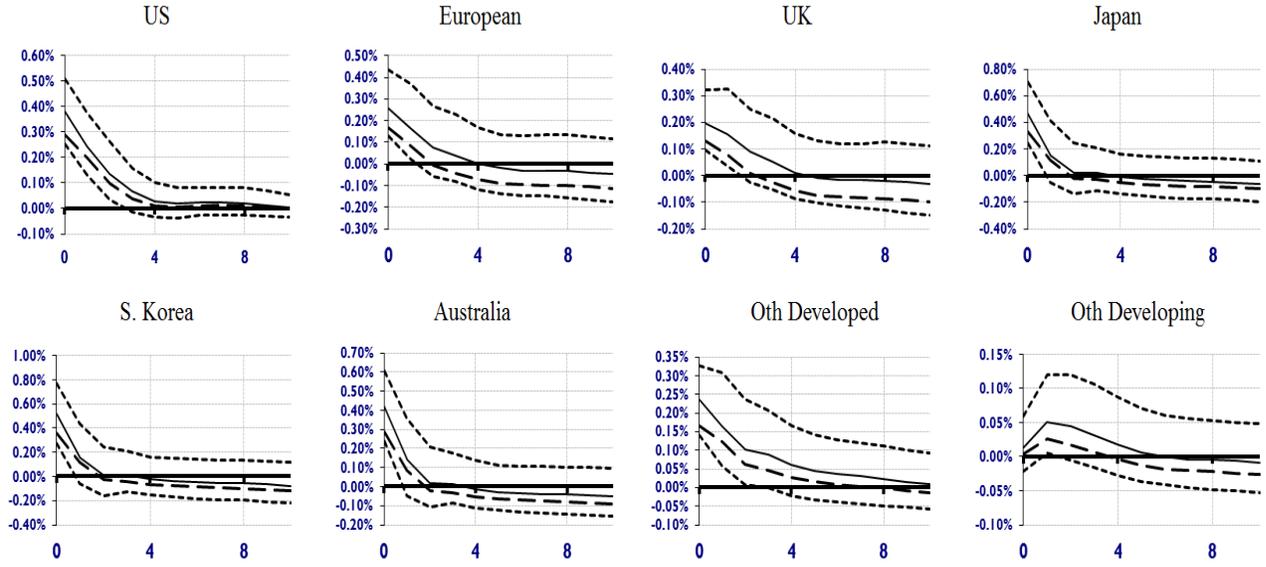
Notes: See Figure A5.

Figure A9: Inflation responses to a composite shock to traded-sector productivity improvement



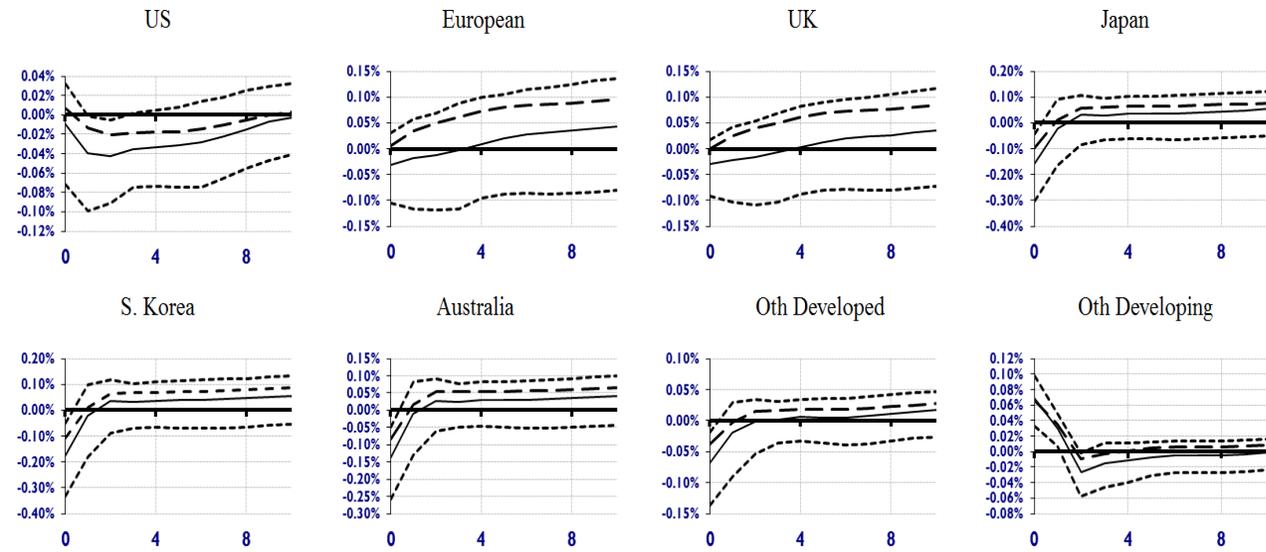
Notes: See Figure A5.

Figure A10: Openness responses to a composite shock to traded-sector productivity improvement



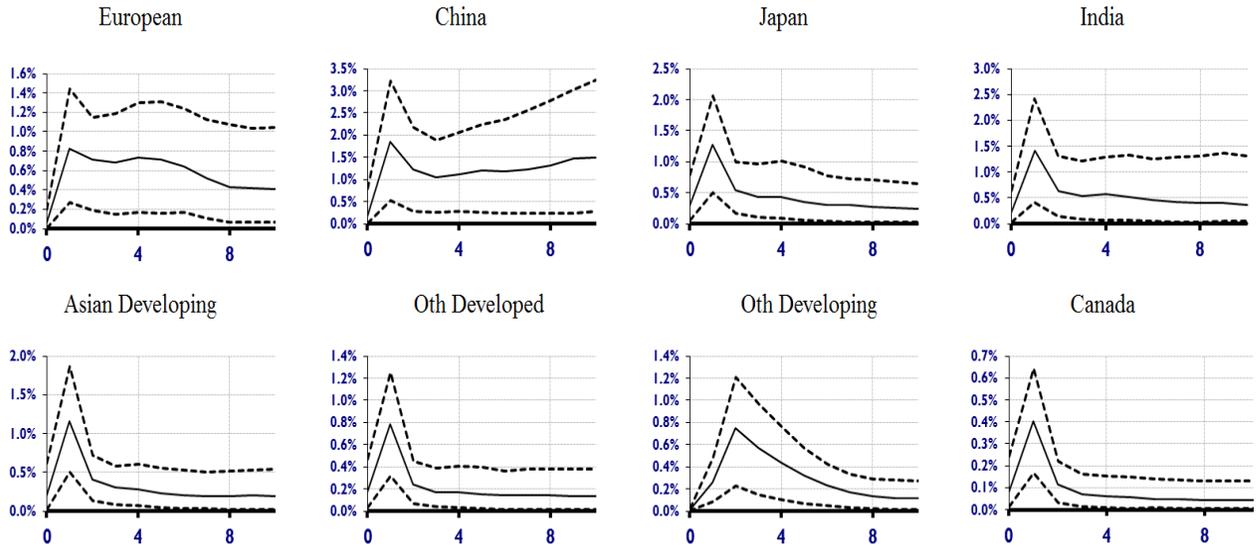
Notes: See Figure A5.

Figure A11: Government consumption share responses to a composite shock to traded-sector productivity improvement



Notes: See Figure A5.

Figure A12: Shares of forecast error variance of real exchange rates explained by a composite shock to traded-sector productivity improvement (%)



Notes: Solid line represents median forecast error variance shares, and dotted lines correspond to 16<sup>th</sup>/84<sup>th</sup> percentile of the posterior distribution. See Figure A5 for the details of countries in each group.

## B Table Appendix

Table B1: List of countries in each group

Country grouping	Members
<b>Developed countries</b>	
European countries ( 9 countries)	Euro Area (Germany, France, Italy, Spain, Netherlands), U.K., Norway, Sweden, Switzerland
Non-European developed countries ( 7 countries)	Japan, South Korea, Singapore, Australia, New Zealand, Canada, USA
<b>Developing countries</b>	
Asian developing countries ( 8 countries)	China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan, Sri Lanka
Non-Asian developing countries ( 6 countries)	Brazil, Mexico, Chile, Turkey, Argentina, South Africa

Notes: We classify countries into developed and developing countries according to the International Monetary Fund's World Economic Outlook Report, April 2010.

Table B2: Contemporaneous effects of foreign variables on their domestic counterparts

Country groups	Domestic dependent variables				
	<i>gov</i>	<i>x</i>	<i>y</i>	$\pi$	<i>si</i>
European countries	0.507** (3.156)	-0.002 (-0.017)	-0.002 (-0.018)	0.227* (1.836)	0.346** (6.234)
Non-European developed countries	0.475** (2.345)	0.113 (0.835)	0.048 (0.518)	0.547** (5.046)	0.453** (3.970)
Asian developing countries	0.499** (2.246)	-0.015 (-0.086)	0.148 (1.426)	0.566** (2.012)	0.388** (2.161)
Non-Asian developing countries	0.119 (0.408)	-0.191* (-1.874)	0.108 (0.475)	-1.703** (-2.786)	0.908** (2.243)
US	0.364 (1.501)	0.550 (1.511)	0.241** (2.617)	0.261** (3.507)	- -

Notes: We specify the VECMX\* for each country  $i$  as:

$$\Delta g_{i,t} = h_{i,0} - \Pi_i z_{i,t-1} - \Phi_{i,2} \Delta g_{i,t-1} + \Psi_{i,0} \Delta g_{i,t}^* + u_{i,t},$$

where  $i = 0, 1, 2, \dots, N$ , and  $t = 1, 2, \dots, T$ . We report the elements of estimated coefficients  $\Psi_{i,0}$  of first differenced contemporaneous foreign variables, which correspond to the elements of their domestic counterparts. For instance, the elasticity of the government consumption share in European countries with respect to the government consumption share in the rest of the world is 0.507. Note that when constructing individual country models, we estimate country models on a country-group basis. Here T-statistics are reported in parentheses. \*,\*\* indicate 10% and 5% significance levels, respectively.

Table B3: F statistics for testing the weak exogeneity of country/region-specific foreign and global variables

Country groups	Foreign variables						
	$x_{i,t}^*$	$y_{i,t}^*$	$gov_{i,t}^*$	$si_{i,t}^*$	$inf_{i,t}^*$	$oil_t$	$\overleftarrow{e}_{i,t}^*$
European countries	1.323 [0.269]	0.025 [0.975]	1.907 [0.152]	2.131 [0.122]	0.243 [0.785]	1.735 [0.181]	
Non-European developed countries	0.820 [0.442]	0.526 [0.592]	0.732 [0.482]	0.842 [0.062]	2.835 [0.062]	1.860 [0.159]	
Asian developing countries	1.254 [0.292]	2.570 [0.056]	0.658 [0.579]	1.773 [0.154]	2.303 [0.078]	1.315 [0.270]	
Non-Asian developing countries	0.291 [0.748]	1.669 [0.192]	0.411 [0.664]	2.457 [0.090]	3.623** [0.030]	2.642 [0.075]	
US	1.421 [0.169]	0.271 [0.979]	0.329 [0.745]	– –	1.077 [0.307]	– –	–0.068 [0.946]

Notes: The weak exogeneity test can be performed by estimating

$$\Delta g_{s,t}^* = c_s + \sum_{r=1}^R \phi_{s,r} ecm_{t-1}^r + \sum_b^B \theta_{s,b} \Delta g_{t-b} + \sum_l^L \theta_{s,l} \Delta g_{t-l}^* + \varepsilon_{s,t},$$

where  $g_{s,t}^*$  is the  $s^{th}$  element of the foreign variable vector  $g_t^*$ ,  $c_s$  is a constant term,  $\varepsilon_{s,t}$  is an error term and  $ecm_{t-1}^r$  are the estimated error correction terms corresponding to the  $r$  cointegrating relationships. We choose the lag orders,  $B$  and  $L$  in the light of serial correlation tests, allowing for a maximum of four lags for both domestic and foreign variables. We test the null hypothesis that the foreign variable is weakly exogenous by testing the joint significance of the estimated error correction terms in the above regression. \*\* denotes statistical significance at the 5% level, and the figures in square brackets are estimated probability values.

Table B4: Forecast error variance shares of real exchange rates explained by the shocks (in percent)

	Productivity shocks				
	h=0	h=1	h=2	h=4	h=8
UK	0.06	0.92	0.72	0.68	0.30
Euro	0.05	0.82	0.71	0.74	0.46
Norw	0.05	0.77	0.66	0.92	0.54
Swed	0.04	0.73	0.63	0.80	0.34
Swit	0.05	0.76	0.70	0.76	0.37
Jap	0.28	1.28	0.54	0.42	0.28
Kor	0.19	0.92	0.28	0.18	0.14
Sing	0.38	1.41	0.38	0.24	0.12
Aus	0.23	1.01	0.35	0.30	0.33
NewZ	0.18	0.86	0.27	0.21	0.21
Cana	0.08	0.40	0.12	0.06	0.04
Chin	0.15	1.85	1.23	1.11	1.32
Mala	0.26	1.40	0.42	0.26	0.17
Indo	0.08	0.71	0.30	0.20	0.15
Phil	0.15	0.98	0.34	0.20	0.13
Thai	0.22	1.26	0.41	0.29	0.25
Indi	0.21	1.42	0.63	0.58	0.41
Paki	0.31	1.78	0.67	0.51	0.26
SriL	0.27	1.59	0.56	0.42	0.24
Braz	0.00	0.17	1.00	0.51	0.13
Mexi	0.00	0.13	0.45	0.24	0.03
Chil	0.02	0.69	1.03	0.63	0.33
Turk	0.01	0.48	0.72	0.43	0.18
Arge	0.00	0.11	0.49	0.55	0.12
S.Afri	0.01	0.60	0.92	0.59	0.29