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Do Policy-Related Shocks Affect Real Exchange Rates of Asian Developing Countries?

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Abstract

This paper examines real exchange rate responses to shocks in exchange rate determinants and monetary policy for eight Asian developing countries. The analysis is based on a panel pseudo-Bayesian structural vector error correction model, and the shocks are identified using sign and zero restrictions. We find that trade liberalization generates permanent depreciation, and higher government consumption causes persistent appreciation. Traded-sector productivity gains induce appreciation but their effects are short-lived. Real exchange rate responses to unexpected monetary tightening are consistent with the Dornbusch overshooting hypothesis and long-run neutrality of monetary policy. The evidence suggests that trade liberalization provides an effective device for driving exchange rate movements.

JEL Classification: C33, C51, E52, F31

Keywords: Real exchange rates, Exchange rate determinants, Vector error correction model, Monetary policy shock, Sign restriction, Penalty function

1 Introduction

Many economists believe that the main causes of the 1997 Asian Financial crisis were dramatic devaluations of Southeast Asian currencies and an inappropriate degree of tightness in monetary policies used for stabilizing their currencies. The collapse of a few Asian foreign exchange markets spread rapidly to put depreciation pressure on other foreign currencies, exacerbated the failure in other financial markets, and finally plunged many Asian regional economies into deep recession. From this situation, we learn that an understanding of real exchange rate behaviour might help us forestall similar future crises and minimize damage that might occur.

This paper aims to examine the roles played by trade liberalization, productivity improvement in the traded sector, contractionary monetary policy and expansionary government consumption in explaining real exchange rate behaviour, and to assess whether these factors have played an

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important role in driving real exchange rate movements in Asia. Our data set comprises annual macroeconomic and financial variables for eight Asian developing countries over the period from 1970 to 2008. We base our analysis on a panel of countries instead of undertaking a country by country analysis, so that panel estimation techniques can improve the efficiency of our parameter estimates. We consider impulse response and variance decomposition.

Our paper makes four main contributions. First, we add to the empirical literature on the relationship between the real exchange rate and its determinants for developing countries, which is currently small because of data limitations. Second, the conventional literature on shock identification in open economy models has relied on structural vector autoregressive (SVAR) models which do not account for long-run relationships between variables. To capture these relationships, we employ an augmented structural vector error correction model (SVECMX*) and apply this to a panel of data. Our model is augmented with foreign variables to capture the influence of foreign countries, particularly main trading partners. We firstly perform tests of long-run equilibrium relationships suggested by theory, and then impose those relationships that are valid on the estimated model. Third, we impose sign restrictions and a zero restriction on impulse responses, and also incorporate a penalty function that reduces the set of admissible impulse responses to a small number of similar estimates. We use a Bayesian approach to implement sign and zero restrictions. This allows for estimation uncertainty in the parameter estimates and deals with nonexact identification of impulse vectors. Fourth, our sectoral productivity differential is constructed using a novel approach for classifying traded and non-traded industries introduced by Dumrongrittikul (2012). This approach allows for different classifications of industries for different countries, and changes in classifications of industries across periods. It also facilitates appropriate study of the Balassa (1964) - Samuelson (1964) (BS) effect.

We find that productivity growth in the traded sector induces currency appreciation on impact as predicted by the BS hypothesis, but that this effect on real exchange rates dies out very rapidly. This is interesting, in light of the recent political pressures that have been placed on fast growing economies in Asia to appreciate their currencies. Other results of impulse response analysis support the most likely paths of real exchange rate reaction suggested by Edwards (1989) and other theoretical models. First, there is strong evidence that trade liberalization significantly generates permanent depreciation and that it also makes a large contribution to real exchange rate fluctuations, especially in the short run. This finding suggests that trade policy should be considered as an effective and powerful instrument for dealing with real exchange rate movements. Second, an increase in government consumption leads to persistent appreciation. Government authorities need to be aware of this when changing a policy on government consumption to influence the economy.

Third, using sign and zero restrictions to identify a monetary policy shock, we overcome exchange rate puzzles i.e. our results are consistent with the Dornbusch's (1976) well-known overshooting hypothesis and long-run neutrality of monetary policy.

The remainder of this paper is organized as follows. Section 2 briefly reviews the existing literature. Section 3 outlines the model and the identification procedure used for recovering the shocks of interest. Section 4 presents the data set, econometric methodology and the empirical analysis. Section 5 concludes.

2 Related Literature

Early work on the relationship between the real exchange rate and its determinants has relied on cross-section comparisons or standard time-series techniques. This can lead to imprecise estimation and inconclusive hypothesis testing when samples are small. More recent studies have turned to panel data cointegration methods. For instance, Chinn (1999) estimates a panel error correction model for fourteen OECD countries and shows that an increase in traded-sector productivity induces a long-run appreciation, while government spending and the terms of trade have no effect on real exchange rates. These results are inconsistent with the work of Galstyan and Lane (2009). They find that government consumption induces a long-run appreciation. Similarly, Lee, Milesi-Ferretti, and Ricci (2008) use panel Dynamic Ordinary Least Squares (DOLS) estimation and focus on forty-eight industrial and emerging economies. Their work suggests that an increase in net foreign assets, the productivity of tradables relative to non-tradables, the commodity terms of trade, the extent of trade restrictions and government consumption cause currency to appreciate.

Much of this literature has focused on developed countries while the empirical evidence for developing countries is quite scant. Moreover, the construction of traded-non-traded productivity differentials is usually based on arbitrary methods; that is, the classification approaches seem unreasonable because they use the same patterns of traded and non-traded sectors among industries for each country. In addition, several existing papers examine the causal effects of exchange rate determinants on real exchange rates by relying only on the study of cointegrating relationships. The shortcoming with this approach is that the presence of a cointegrating relationship does not provide information on the direction of causality.

In this paper, we focus on real exchange rate behaviour in Asian developing countries. An estimation issue related to the data limitations of developing countries is solved by using panel data methods. For the construction of sectoral productivity, we use a novel classification approach introduced by Dumrongrittikul (2012) that allows for country-specific heterogeneity over each industry and changes in classification across time. Further, we extend previous studies that mostly rely on estimated cointegrating relationships by using impulse response analysis to examine the causal effects of exchange rate determinants on real exchange rates in the short run and long run.

3 The Model and Identification Procedure

Our model relates to eight Asian developing countries - China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan and Sri Lanka - and it contains domestic and foreign variables. Instead of using US variables as a proxy for foreign variables, we consider twenty-six countries/regions when constructing country-specific foreign variables.[1] 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 of country-specific foreign variables and a global factor (i.e. oil prices). Each country-specific foreign variable is constructed as a country-specific weighted average of corresponding domestic variables of all other countries/regions:

$$g^*_{i,s,t} = \sum_{j=1}^{26} w_{i,s,j} g_{j,s,t}, \quad i = 1, 2, ..., N,$$

where N $(= 8)$ is the number of countries in the model and $g_{i,s,t}(g^*_{i,s,t})$ is the element of $g_{i,t}(g^*_{i,t})$ corresponding to variable s. Following the literature, we use a weighting scheme based on bilateral trade exposure (average trade shares over the period 2002-2008) to capture the relative importance of country j to country i. In particular, $w_{i,s,j}$ is the trade share of country j in the total trade (imports + exports) of country i with all of its trade partners associated with variable s, such that $w_{i,s,i} = 0$ and $\sum_{j=1}^{26} w_{i,s,j} = 1$.[2]

3.1 Structural Vector Autoregressive Model

We construct a panel model for the group of Asian developing countries because the spans of data are short and the use of panel data methods will improve the efficiency of our parameter estimates. The key assumption for the purposes of estimation and inference is that foreign and global variables are weakly exogenous, compatible with a limited degree of weak dependence across idiosyncratic shocks. To satisfy this property, we assume that all economies in the model are small relative to the world economy. This is a reasonable assumption given that our country group consists of eight

[1] In addition to eight Asian developing countries, our set of countries comprises the United Kingdom, Euro Area (Germany, France, Italy, Spain, Netherlands), Norway, Sweden, Switzerland, Australia, New Zealand, Canada, the United State, Korea, Japan, Singapore, Brazil, Mexico, Chile, Argentina, South Africa and Turkey. We use a superscript * to denote a foreign variable and no superscript to indicate a domestic variable. The oil price, oil_t, is a global variable that has no superscript.

[2] Data sources for each series and a 26×26 matrix of the trade shares used for constructing the country-specific foreign variables are available upon request.

Asian developing countries. We perform formal tests of weak exogeneity of foreign variables along the lines described in Johansen (1992), and our test results support this assumption.

We use residual serial correlation test results to choose an augmented vector autoregressive (VARX*) model, with a third-order dynamic specification for domestic variables and a first-order dynamic specification for foreign variables. The resulting VARX*(3,1) model can be written as

$$g_{i,t} = \Phi_1 g_{i,t-1} + \Phi_2 g_{i,t-2} + \Phi_3 g_{i,t-3} + \Psi_0 g^*_{i,t} + \Psi_1 g^*_{i,t-1} + u_{i,t}, \tag{1}$$

for $t = 1, 2, ..., T$ and $i = 1, 2, ..., N$. The notation is such that Φ_1, Φ_2 and Φ_3 are $k_i \times k_i$ matrices of coefficients associated with lagged endogenous/domestic variables, Ψ_0 and Ψ_1 are $k_i \times k_i^*$ matrices of coefficients associated with weakly exogenous/foreign variables, and $u_{i,t}$ is a $k_i \times 1$ vector of reduced-form residuals with a variance-covariance matrix $\Sigma = E[u_{i,t}u'_{i,t}]$ for all i and all t. Country-specific fixed effects are allowed in our model via the inclusion of country-specific dummy variables. However, we have dropped the intercept and country-specific dummy variables from equation (1) to simplify the notation.

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

$$\Delta g_{i,t} = -\Pi z_{i,t-1} - (\Phi_2 + \Phi_3)\Delta g_{i,t-1} - \Phi_3 \Delta g_{i,t-2} + \Psi_0 \Delta g^*_{i,t} + u_{i,t}, \tag{2}$$

where

$$\Pi = (I - \Phi_1 - \Phi_2 - \Phi_3, -\Psi_0 - \Psi_1) \text{ and } z_{i,t-1} = \left(g'_{i,t-1}, g^{*'}_{i,t-1}\right)'.$$

It is easy to see that the cointegrating relationship among variables is summarized in a $k_i \times (k_i + k_i^*)$ matrix Π. Suppose that the rank of Π is $r_i \leq k_i$, implying that there are r_i long-run relationships among the variables. The matrix $\Pi = \alpha\beta'$, where α is a $k_i \times r_i$ loading matrix of full column rank and β is a $(k_i + k_i^*) \times r_i$ matrix of cointegrating vectors of rank r_i. We impose cointegrating restrictions $\widehat{\beta}$ computed by using panel DOLS, when estimating the VECMX*. Thus we can rewrite equation (2) as

$$\Delta g_{i,t} = B_0 ecm_{i,t-1} + B_1 \Delta g_{i,t-1} + B_2 \Delta g_{i,t-2} + B_3 \Delta g^*_{i,t} + u_{i,t}, \tag{3}$$

where $ecm_{i,t-1} = \widehat{\beta}' z_{i,t-1}$, $B_0 = -\alpha$, $B_1 = -(\Phi_2 + \Phi_3)$, $B_2 = -\Phi_3$ and $B_3 = \Psi_0$.

Our interest is to examine impulse responses to economically meaningful structural shocks $v_{i,t}$ associated with the VARX* in (1). This leads to the problem of identification, given the correlation among reduced-form residuals $u_{i,t}$. Recall that the dimension of $g_{i,t}$ is k_i. We adopt an appealing assumption in the VAR literature that there are k_i fundamental innovations $v_{i,t}$, which are mutually independent and normalized to have a variance of 1, i.e. $E(v_{i,t}v'_{i,t}) = I_{k_i}$. We require a identifying matrix A such that $u_{i,t} = Av_{i,t}$ to obtain independence of the fundamental innovations. Note that the immediate impact or impulse vector of the j^{th} structural innovation (which is the j^{th} element of the vector $v_{i,t}$) of a one standard deviation shock in $v_{i,t}$ on each endogenous variable in the system can be represented by the j^{th} column of the matrix A, a_j.

3.2 Sign Restrictions with a Penalty-Function Approach

Traditional identifications are commonly based on zero contemporaneous and/or long-run restrictions. The results from these approaches depend on the chosen decomposition of a variance-covariance matrix. This leads to the imposition of very stringent restrictions, most of which may not rely on theoretical considerations. In this paper, we avoid these problems by imposing sign restrictions on the impulse response functions to identify four types of underlying disturbances; trade liberalization, productivity growth, monetary policy and government consumption shocks. This new identification strategy was developed by Uhlig (2005) and extended by Mountford and Uhlig (2009). This approach requires only a set of economically plausible restrictions that are often used implicitly by researchers. It makes a priori theoretical restrictions explicit and leaves the question of interest open. Underlying shocks can be identified by examining whether the signs of the corresponding impulse responses are accepted by a priori consensus considerations. The imposition of sign restrictions leads to results that are robust to reordering variables and selecting a particular Cholesky decomposition.

As shown by Uhlig (2005), identification does not depend on any particular matrix A. If there exists a k_i-dimensional vector m of unit length such that $a = \overrightarrow{A}m$, where $\overrightarrow{A}\overrightarrow{A}' = \Sigma$ and $\overrightarrow{A}$ is any arbitrary decomposition of Σ such as a lower triangular Cholesky factor,[3] we can obtain an impulse vector a even though the true matrix A is not identified. Uhlig uses this property to show that the impulse response $r_a(h)$ at horizon h to the impulse vector a can be computed as a linear combination of the impulse responses obtained using the Cholesky decomposition of Σ. This can be represented as

$$r_a(h) = \sum_{j=1}^{k_i} m_j r_j^c(h),$$

where m_j is the j^{th} element of m; $r_j^c(h) \in \mathbb{R}^{k_i}$ is a $k_i \times 1$ vector of the impulse response at horizon h to the j^{th} shock in a Cholesky decomposition of Σ, i.e. the j^{th} column of $\overrightarrow{A}$. As our study focuses on four underlying shocks, we need to characterize an impulse matrix $[a^{(1)}, a^{(2)}, a^{(3)}, a^{(4)}]$ of rank

[3] The Cholesky decomposition is only used as a computational tool when implementing the sign restriction approach, but it is not used for the purpose of identification.

4, rather than all impulse vectors in the matrix A.

Moreover, it is interesting to perform variance decompositions or in other words, compute how much a shock contributes to the variance of the h-step ahead forecast error. The fraction $\phi_{a,s}(h)$ of the variance of the h-step ahead forecast revision for variable s in response to an impulse vector a can be obtained by

$$\phi_{a,s}(h) = \frac{(r_{a,s}(h))^2}{\sum_{j=1}^{k_i}(r_{j,s}^c(h))^2},$$

where the additional index s picks the response corresponding to variable s.

Following Uhlig (2005), we deal with the sampling uncertainty of the OLS estimates and the nonexact identification of impulse matrices by using a Bayesian method to implement the sign restrictions. A Monte Carlo integration is performed. Given the estimated VECMX*, we take a joint draw from the posterior of the Normal-Wishart distribution for (B, Σ) and a draw from a uniform distribution over the unit sphere for candidate m vectors. The Cholesky decomposition factor, $\overrightarrow{A}$ is computed using a draw of Σ from the posterior. Consequently, we can calculate the candidate impulse vector as $a = \overrightarrow{A}m$.

We apply a sign restriction approach with a penalty function, rather than a pure-sign restriction approach. The main difference between these two approaches is that with a pure-sign restriction approach, all impulse vectors satisfying the sign restrictions are considered equally for determining the impulse responses, while the penalty-function approach chooses the best of all impulse vectors for each draw of (B, Σ) via minimizing a criterion function. Although no impulse response might satisfy all sign restrictions, the impulse vector which generates responses that satisfy the sign restrictions as closely as possible is considered. Thus, with the penalty-function approach, it is possible to obtain impulse response functions with smaller standard errors.

Let l_{r+} be the set of variables for which the impulse response is restricted to be positive and l_{r-} be the set of variables for which the impulse response is restricted to be negative. H_{re} is the last period that responses are constrained. The standard deviation of the first-differenced variable s, denoted by σ_s is used for rescaling impulse responses, or in other words, generating standardised impulse responses so that the deviations across different impulse responses are comparable to each other. To implement the penalty-function approach, we minimize the criterion function $\Psi(a)$ in order to find the best impulse vector a for each draw of (B, Σ). That is, we solve

$$a = \arg \min_{a=\overrightarrow{A}m} \Psi(a)$$

for

$$\Psi(a) = \sum_{s \in l_{r+}} \sum_{h=0}^{H_{re}} ff(-\frac{r_{a,s}(h)}{\sigma_s}) + \sum_{s \in l_{r-}} \sum_{h=0}^{H_{re}} ff(\frac{r_{a,s}(h)}{\sigma_s}), \tag{4}$$

where the penalty function suggested by Uhlig (2005) is

$$ff(ww) = \begin{cases} ww & \text{if} \quad ww < 0, \\ 100 \times ww & \text{if} \quad ww \geqslant 0. \end{cases} \tag{5}$$

It is obvious that the penalty function is asymmetric when imposing sign restrictions, i.e. we penalize wrong responses 100 times more than we reward correct responses. Using numerical minimization on the criterion function (4), we can identify the first shock $a^{(1)} = \overrightarrow{A}m^{(1)}$.

We add the restriction that the second shock is orthogonal to the first shock to identify the second shock. Moreover, following Mountford and Uhlig (2009), we can easily impose a zero contemporaneous restriction on the impulse response of variable s by imposing a restriction on the vector m such that $Rm = 0$, where $R = [r^c_{1,s}(0), ..., r^c_{k_i,s}(0)]$. Therefore, additionally imposing orthogonality conditions and a zero restriction, we minimize the problem below:

$$a = \arg \min_{a=\overrightarrow{A}m, Rm=0, m'm^{(1)}=0} \Psi(a).$$

3.3 Identifying Assumptions and Implementation Based on Sign Restrictions

Table 1 summarizes the set of restrictions adopted in this paper. In Table 1, q represents the real exchange rate, x is the traded-non-traded productivity differential, y is real GDP, gov is the government consumption share, $open$ is the degree of openness in the economy, si is the nominal short-term interest rate, and π is the inflation rate.

Table 1: Identifying restrictions

	q	x	y	gov	$open$	si	π
Trade liberalization shock					+		−
Productivity improvement shock		+	+				
Contractionary monetary policy shock		0	−			+	−
Government consumption shock			+	+			+

Notes: 1) + (-) means positive (negative) response of the variables in columns to shocks in rows. 0 means no response. 2) The sign restrictions are imposed from impact to lag 1, while a zero restriction is imposed on impact only.

We identify a trade liberalization shock as an unexpected rise in the international trade share for a year. The response of inflation is restricted to decrease since trade liberalization is normally viewed as a reduction in import tariff. This will lower prices in the domestic country. A productivity shock is identified as a shock that causes productivity in traded sectors relative to non-traded sectors and real output to increase for a year. These restrictions correspond to the BS hypothesis which expects that productivity improvement in traded sectors is more rapid in countries with higher growth rates than those with lower ones. The government consumption shock is characterised by a rise in government consumption, real GDP and inflation for a year.

We achieve the identification of a contractionary monetary policy shock by imposing a mixture of sign and zero restrictions. Following a contractionary monetary policy shock driven by an increase in the interest rate, there is a drop in inflation and real output over a year, whereas the productivity differential between traded and non-traded sectors is assumed to be initially unchanged. The zero restriction on the impact response of the productivity differential between traded and non-traded sectors is plausible because an unexpected change in monetary policy should influence the productivity in both sectors in the same way on impact. Moreover, although the conditions on inflation and output can be controversial, i.e. observed price puzzles or output puzzles are not easily explained, the sign restrictions imposed here are consistent with the standard New Keynesian model and the empirical works e.g. Farrant and Peersman (2006).

In short, these restrictions are in line with the theoretical and empirical literature and are sufficient to uniquely disentangle the shocks of interest. We do not impose any responses of uncertain sign. As our focus is on the responses of the real exchange rate, we leave all its responses unrestricted. We firstly identify a trade liberalization shock and then identify an orthogonal productivity shock, an orthogonal monetary policy shock, and an orthogonal government consumption shock, in that order. The idea behind this ordering is that it is difficult to distinguish the movement of each variable caused by a shock to that variable from the contemporaneous movement in that variable caused by other shocks. The orthogonality condition can help to identify the shock by filtering out the contemporaneous responses of each variable to other shocks. We decide to begin with the trade liberalization shock because international trade is normally slow to respond to a shock, given the sluggish behavior of real economic activity between countries. Then we choose the shock to traded-sector productivity as a second shock for identification because the productivity will take time to adjust, similar to the international trade share. Nonetheless we have checked that our results are robust to change in the order of these first two shocks (See Table B2 in Appendix B). For monetary policy and government consumption shocks, we choose to order these shocks after trade liberalization and productivity shocks in order to filter out the effects of the latter shocks.

For computation, we find a "best impulse matrix" by undertaking the numerical minimization of the above criterion function $\Psi(a)$ in equation (4) on the unit sphere, given each draw of B and Σ. We parameterize the space of unit-length vectors by using stereo projection,[4] which is available

[4] The stereo projection is a way of drawing the unit sphere onto the plane through the equator. With this technique, the angles at which curve cross each other are preserved but areas and distances are distorted.

in the RATS statistical package. We do the minimization procedure twice for each draw, starting it from two different initial random vectors in order to check whether the best impulse vector we obtain is the optimal solution. In particular, we examine whether the two minima found are very close or the same. If they are the same or different by less than 0.01, we keep the impulse vector. In contrast, if they generate total values of the penalty that differ by more than 0.01, we keep only the vector which generates the smaller value of the total penalty, and we discard the other. Therefore given each draw of B and Σ, we will obtain a selected impulse matrix for computing impulse responses. Then we draw a new B and Σ, and start a new minimization procedure using the last set of minimizers as one of initial vectors. We continue and repeat these procedures until we have acquired 1,000 draws of B and Σ, generating 1,000 best impulse matrices and a sample of 1,000 impulse responses. Given this sample, we find the impulse responses at the 16^{th}, 50^{th} and 84^{th} percent quantiles for each step-ahead forecast.

4 Empirical Investigation

4.1 Data Description

We employ a panel data set that includes annual time series from 1970 to 2008.[5] The data set covers eight developing countries in Asia - China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan and Sri Lanka. Following Dees et al. (2007), the log real effective exchange rate is defined as $q_{i,t} = (e_{i,t} - p_{i,t}) - (e^*_{i,t} - p^*_{i,t})$, where $e_{i,t}$ is the log nominal exchange rate with respect to the US and $p_{i,t}$ is the log CPI for country i during the period t. By construction, an increase in the real exchange rate represents a real *depreciation*. Given theoretical models of real exchange rate determination and data availability, our set of real exchange rate fundamentals include the log traded-non-traded productivity differential ($x_{i,t}$);[6] log real GDP ($y_{i,t}$); the log terms of trade ($tt_{i,t}$), defined as the ratio of the export price index to the import price index; the log government consumption share ($gov_{i,t}$), measured as the ratio of government consumption to GDP; and the log openness of the economy ($open_{i,t}$), measured as the ratio of the sum of exports and imports to GDP.

When constructing the sectoral productivity differential ($x_{i,t}$), it is important to classify economic activities into traded and non-traded sectors. This is the weakest aspect of previous studies of the BS hypothesis. Therefore, in this paper we use a new approach introduced by Dumrongrit-

[5] The length of each series varies slightly from country to country.

[6] Labour productivity is used as a proxy variable for total productivity due to data limitations. Productivity in each sector is computed as the ratio of value added in constant local currency unit to the number of employees in each sector.

tikul (2012). This approach allows for different patterns among industries in different countries and for trade endogeneity. It is based on two main concepts: tradability and sectoral price comovement. The first concept is that traded goods and services are normally traded in international markets, while goods and services in the non-traded sectors can be traded only in domestic markets. The classification approach uses international trade information from Input-Output tables to assess the tradability of each industry. The second concept is that the price of traded goods and services is more likely to follow the Purchasing Power Parity (PPP) and the Law of One Price than the price of non-traded goods and services. This can be examined by using econometric methods to test the comovement of prices in each industry with world prices.

Edwards (1989) shows that although the long-term real exchange rate relies on real variables only, both nominal and real variables can influence the short-term real exchange rate. Therefore in our empirical model, we also include the nominal short-term 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 inflation rate ($\pi_{i,t} = p_{i,t} - p_{i,t-1}$) and the log oil price index (oil_t) that accounts for global unobserved factors. Data sources are provided in Appendix A.

Each variable is tested for the presence of a unit root by using both time-series and panel unit root tests. The results show that all of the domestic and foreign variables in levels and all of the differences between domestic and foreign variables are approximately $I(1)$, except interest rate differentials ($si_{i,t} - si^*_{i,t}$) and domestic inflation ($\pi_{i,t}$).

4.2 Testing for the Long-Run Relationships

We pay careful attention to testing long-run relationships to avoid a misspecification problem. We use a parsimonious approach, instead of the traditional system approach, because our system has large dimensionality and the performance of the traditional approach is generally very poor in this setting. We conduct tests for possible long-run relationships which are borrowed from economic theory as possible candidates:

Purchasing Power Parity	$q_{i,t}$	$\sim$	$I(0)$	(1)
Fisher Equation	$si_{i,t} - \pi_{i,t}$	$\sim$	$I(0)$	(2)
Output Convergence	$y_{i,t} - y^*_{i,t}$	$\sim$	$I(0)$	(3)
Uncovered Interest Parity	$si_{i,t} - si^*_{i,t} - E(\Delta e_{i,t+1})$	$\sim$	$I(0)$	(4)

and a hybrid Balassa-Samuelson (1964) and Edwards (1989) model

$$q_{i,t} - \lambda_1(x^*_{i,t} - x_{i,t}) - \lambda_2(y^*_{i,t} - y_{i,t}) - \lambda_3(gov^*_{i,t} - gov_{i,t}) - \lambda_4(tt_{i,t}) - \lambda_5(open_{i,t}) \sim I(0) \qquad (5)$$

The first relationship, called PPP is the well-known theory of long-term equilibrium exchange rates based on the relative price levels between countries. The second relationship is the Fisher Equation which shows the relationship between nominal and real interest rates under inflation. Note that the results of the unit root tests suggest that $\pi_{i,t}$ is stationary, so this relationship is reduced to $si_{i,t} \sim I(0)$. The third relationship represents the relative output convergence condition loosely derived from the Solow-Swan neoclassical growth model.[7] The fourth relationship is the uncovered interest parity (UIP) condition which relates the difference between domestic and foreign nominal interest rates to the expected future change in the exchange rate. Since the results of the unit root tests show that $E(\Delta e_{i,t+1})$ is $I(0)$, this relationship can be reduced to $si_{i,t} - si^*_{i,t} \sim I(0)$. The fifth relationship describes the relationship between exchange rates and its determinants, and as the movement of real exchange rates depends not only on domestic impacts but also on external impacts from outside countries, we choose to use variables in relative terms. However, we do not use relative terms for the terms of trade and open variables because they have already accounted for the interaction between domestic and foreign countries by construction.

We conduct panel unit root tests on $q_{i,t}$ and $si_{i,t}$ to check the validity of the relationships (1)-(2). The results show that $q_{i,t}$ and $si_{i,t}$ are approximately $I(1)$, suggesting that PPP and the Fisher Equation do not hold.

In order to test for the long-run relationships (3)-(5), we apply four residual-based cointegration tests suggested by Pedroni (1999).[8] The tests are based on the null hypothesis that for each country in the panel the variables of interest are not cointegrated, while the alternative hypothesis is that there exists a single cointegrating vector for each country in panel. These approaches allow cointegrating vectors to be different for each country. Since the tests are based on the assumption of cross-sectional independence in the error term, we include a set of common time dummies in the hypothesized cointegrating regression to accommodate some forms of cross-sectional dependence across different countries.

[7] The Solow-Swan model actually shows convergence of output per capital, not output.

[8] We consider four test statistics, instead of all seven test statistics because Pedroni (2004) shows that in a situation similar to ours, panel-t statistics and group-t statistics have higher power than other test statistics.

Table 2: Cointegration tests for the relationships (3)-(5)

Variables	Panel PP	Group PP	Panel ADF	Group ADF
$y_{i,t}, y^*_{i,t}$	0.32	0.48	0.03	0.03
$si_{i,t}, si^*_{i,t}$	0.00	0.00	0.00	0.00
$q_{i,t}, (x^*_{i,t} - x_{i,t}), (y^*_{i,t} - y_{i,t}),$ $(gov^*_{i,t} - gov_{i,t}), tt_{i,t}, open_{i,t}$	0.86	0.94	0.88	0.85
$q_{i,t}, (y^*_{i,t} - y_{i,t}), open_{i,t},$ $(gov^*_{i,t} - gov_{i,t})$	0.08	0.09	0.01	0.00

Notes: 1) The number in the table reports the P-value for the tests under the null hypothesis of no cointegration.
2) The lag length in ADF-type regression used in Pedroni's tests is selected by AIC with a maximum of 4 lags.

Table 2 reports the results of the panel and group t-statistics of Pedroni's cointegrating tests. Two of Pedroni's tests suggest that output convergence does not hold, corresponding to the results of unit root tests on $(y^*_{i,t} - y_{i,t})$. However, we find strong evidence for UIP. This finding is consistent with what we expect, as nowadays financial markets in most countries are integrated with each other, and have become more like a global financial market. The evidence does not support relationship (5). However, as (5) relates to a large set of variables and the span of our data set in each country is short, the panel test statistics might have low power and poor performance. For this reason, we test all possible subsets of the real exchange rate and its fundamentals by dropping one fundamental at a time, stopping if a cointegrating relationship is found. We find that there is a cointegrating relationship between the real exchange rate and its three fundamentals i.e. $(y^*_{i,t} - y_{i,t}), (gov^*_{i,t} - gov_{i,t}), open_{i,t}$. See Table 2.

4.3 Estimating and Interpreting the Cointegrating Vector

We use within-dimension panel DOLS estimation developed by Mark and Sul (2003) to estimate the cointegrating vector.[9] The within-dimension estimators are somewhat restrictive since cointegrating vectors are supposed to be homogeneous across cross-sectional units. Nevertheless, given that our country set includes country members with similar features, allowing for heterogeneity across countries through heterogeneous short-run dynamics and country-specific fixed effects seems sufficient to capture the heterogeneity in the country members of the panel. We can also obtain more precise point estimates of the cointegrating vectors due to the improvement of finite-sample estimation by using the panel. The estimated DOLS equation for the real exchange rate relationship is

$$\widehat{q}_{i,t} = d_i + \theta_t - \underset{(0.074)}{0.539}(y^*_{i,t} - y_{i,t}) + \underset{(0.097)}{0.276}(gov^*_{i,t} - gov_{i,t}) + \underset{(0.056)}{0.071} open_{i,t} + \delta_i' sd_{i,t}, \quad (6)$$

[9] This is because the DOLS estimator has smaller size distortions and outperforms the OLS and the fully modified OLS estimators in both finite and infinite samples, according to the findings of Kao and Chiang (2000).

where $i = 1, 2, ..., N$; $sd_{i,t} = (\Delta fd'_{i,t-1}, \Delta fd'_{i,t}, \Delta fd'_{i,t+1})'$ where $fd_{i,t} = ((y^*_{i,t} - y_{i,t}), (gov^*_{i,t} - gov_{i,t}), open_{i,t})'$; d_i is a country-specific effect; θ_t is a common time-specific factor which is used to capture some forms of cross-sectional dependence across countries. The coefficients of real exchange rate fundamentals capture their long-run impact on the real exchange rate. Note that standard errors of coefficient estimates are reported in parentheses.

The signs of the estimates correspond to our expectation i.e. after controlling for other factors, an increase in the degree of openness in the economy is associated with a real depreciation and expansionary government consumption is associated with a real appreciation. However, the estimate of economic growth has an unexpected sign, suggesting that rapid economic growth is associated with a real depreciation, after controlling for the other two factors. A possible explanation is that economic growth in these countries may be mainly due to expansion in non-traded sectors. This means that the price of non-traded goods may decrease. A decline in the price of non-traded goods implies a real depreciation as the real exchange rate of a small open economy depends only on the domestic price of non-traded goods.

4.4 Impulse Response Analysis

We compute impulse response functions to determine how quickly shocks to these fundamentals affect real exchange rates and other variables, and how large these impacts are. The VECMX* in (2) comprises seven endogenous variables and six weakly exogenous variables as follows:[10]

$$\begin{aligned} g_{i,t} &= [q_{i,t}, x_{i,t}, y_{i,t}, gov_{i,t}, open_{i,t}, si_{i,t}, \pi_{i,t}]' \text{ and} \\ g^*_{i,t} &= [x^*_{i,t}, y^*_{i,t}, gov^*_{i,t}, si^*_{i,t}, \pi^*_{i,t}, oil_t]'. \end{aligned}$$

Note that when we examine long-run relationship (5), most variables are in differentials (differences between domestic and foreign variables); however, we separate them out into domestic variables and foreign variables when estimating the model.[11]

We mainly focus on the responses of the real exchange rate to four structural innovations - trade liberalization, productivity growth, contractionary monetary policy and expansionary government consumption shocks. Figure 1 shows the median impulse responses of the real exchange rate together with the 16^{th} and 84^{th} percentile responses over nine years after the shocks. We also use

[10] Note that we have dropped the terms of trade ($tt_{i,t}$) from the set of variables. In addition to the consideration of parsimony, this is because our results show that the terms of trade has no long-run relationship with real exchange rates. Also previous empirical work, e.g. Li (2004), suggests that the terms of trade is insignificant for determining the short-run movement of real exchange rates.

[11] We find weak evidence of error cross-sectional dependence when we do this. However, it does not have any significant effect on the responses of the real exchange rate to the four shocks.

Fry and Pagan's (2011) approach to find a single model whose impulse responses are as close to the median responses as possible; however, it does not produce any significant different results from the median responses as shown in Figure 1.[12] Recall that the exchange rate is defined so that it declines as the value of home country currency increases. Note that the impulse responses of other variables are provided in Figures C1-C4, Appendix C.

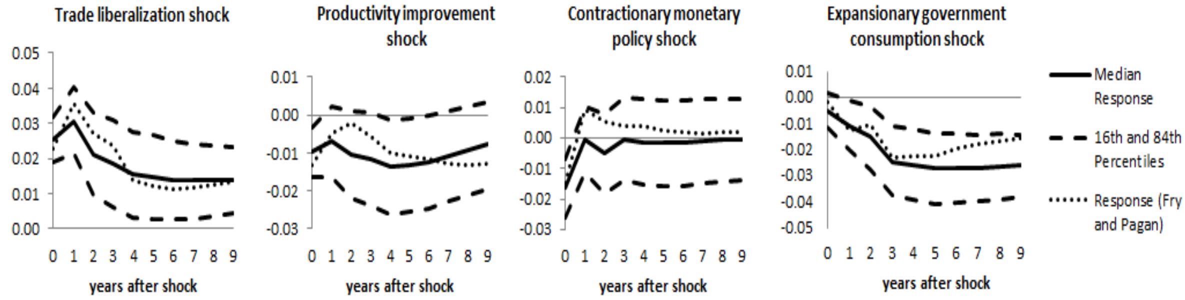

Figure 1: Impulse responses of real exchange rates to each shock of one standard deviation in size.

We base most of our analysis on the Edwards (1989) model because it is used extensively in the empirical literature of developing countries. Edwards explains the reactions of the real exchange rate to a change in its fundamentals using an intertemporal general equilibrium model of a small open economy that can account for all essential aspects of the question we are addressing. According to his model, the real exchange rate depends only on the domestic price of non-traded goods.

4.4.1 Trade Liberalization Shock

The effect of trade liberalization is unclear in the Edwards (1989) model, and it depends on the assumptions imposed. Edwards supposes that there is a permanent decline in import tariffs. This will lead to a positive income effect, implying a rise in the demand for all goods and their prices. Thus, the income effect results in a real appreciation. In contrast, when a small economy liberalizes its trade, demand for importables increases. If all goods are substitutes in consumption, a fall in import prices will reduce the demand and price of non-tradables. This in turn entails a real depreciation. Therefore, the long-run response of the real exchange rate to trade liberalization depends on whether the substitution effect or income effect dominates.

As shown in Figure 1, our analysis provides evidence that trade liberalization results in a long-run real depreciation of the domestic currencies, suggesting that the substitution effect dominates the positive income effect. This is the type of reaction of the equilibrium real exchange rate that,

[12] We are aware of the fact that China's economy might be different from the other countries in panel. Thus, as a robustness check, we re-estimate the model in this country group without China and examine the resulting impulse response analysis. Results remain broadly unaffected. Further, according to the existing literature e.g. Granville and Mallick (2009), Uhlig's methodology is robust to non-stationarity of series including breaks, and therefore it is not necessary to include any dummy variables to account for dramatic regime changes or the Asian financial crisis.

according to Edwards and a variety of theoretical models,[13] is the most likely to occur.

Figure C1 shows that the trade liberalization shock also causes a rise of real output, which might arise because a trade reform can have a positive effect on the expected future return. This provides an incentive for domestic firms to compete or seek new markets overseas, and encourages them to invest more and also attract new investment in the country. Further, the domestic country benefits from gains in trade, with more efficient allocation of resources being distributed across the production of different goods (comparative advantage). Such gains will very likely lead to productivity growth in the non-traded sector. This raises real GDP in the long run.

4.4.2 Traded-Sector Productivity Improvement Shock

The BS hypothesis provides an explanation of long-run real exchange rate behavior that is based on productivity differentials between traded and non-traded goods. The explanation is that higher productivity growth in the traded sectors (given that productivity growth in the non-traded sector for all countries is slow) induces a rise in wages in the traded sectors, and this raises wages in the non-traded sectors due to labour mobility. The non-traded sectors need to raise their prices to maintain profit margins. The law of one price in traded sector goods then implies that a higher price of non-traded goods at home than abroard will induce a long-run appreciation of the domestic currency. Although our result shows that a productivity shock causes a contemporaneous appreciation as predicted by the BS hypothesis, we also find that its effect dies out in the short run. This is consistent with observed exchange rates in rapidly growing countries in Asia such as China and India, since their currencies have not experienced long-run appreciation.

Our results differ from the traditional BS model and the related empirical literature for two main reasons. The first reason is that according to the BS hypothesis, the traded-non-traded productivity differential is only one factor that can cause a persistent deviation of the real exchange rate from its equilibrium. In our model we examine the effect of a productivity gain after controlling for other factors that may have influenced the real exchange rate. The effect of productivity differentials on the real exchange rate may be lowered after capturing these factors, or in other words, the estimated productivity effect based on a two-variable model may pick up the effects of other factors on the behaviour of the real exchange rate.

The second reason is that various measures have been used as a proxy for sectoral productivity differentials. One common shortcoming of these measures is that the classification of the traded and

[13] Dornbusch (1974) shows that if nontradables are substitutes for tradables, a reduction in import tariffs will lead to an equilibrium real depreciation. Similarly, Khan and Ostry (1992) show that a tariff reduction will result in a real depreciation, assuming that all goods are normal and the substitution effect of relative price changes dominates the income effect.

non-traded sectors in all countries is the same and does not change across time. Edwards (1989) mentioned this in his empirical work after he found a result that contradicted the BS hypothesis. As noted above, we have used the classifications introduced by Dumrongrittikul (2012), which change across countries and periods. Nevertheless, we still find that a productivity differential shock only causes a short-run appreciation.[14] This might be because some of the BS assumptions are not satisfied, particularly the labour arbitrage condition. That is, in most Asian developing countries, industrial productivity seems to be higher than rural productivity. However, rural workers are more likely to move to non-traded service sectors than industrial sectors (that are typically traded). Therefore, traded-sector productivity gains need not raise wages and prices in non-traded sectors.

4.4.3 Contractionary Monetary Policy Shock

Standard theory suggests that an unanticipated increase in domestic interest rates causes net inflows on the capital account, boosting the supply of foreign currencies. As a result, the price of foreign currency falls and this leads to an impact appreciation of the domestic currency. If UIP holds, a depreciation of the domestic currency in the future is expected, due to an increase in domestic interest rates relative to foreign interest rates. The Dornbusch overshooting hypothesis follows these predictions and suggests that ceteris paribus, an unanticipated tightening in monetary policy leads to an impact appreciation beyond its long-run value, and then this is followed by a depreciation towards the terminal value implied by UIP.

The price puzzle and the output puzzle do not occur in our study by construction. We overcome exchange rate puzzles found in past empirical research on monetary policy (e.g. Scholl and Uhlig (2008)). In particular, our results find evidence in favour of the Dornbusch model, i.e. a contractionary monetary policy shock causes the domestic currency to appreciate on impact and then depreciate back rapidly to baseline.

In addition to having no long-run response of real exchange rates to the shock, a tightening in monetary policy has only short-run effects on other real variables i.e. real GDP and productivity, consistent with the New Keynesian view of long-run neutrality of monetary policy (See Figure C3).

4.4.4 Expansionary Government Consumption Shock

Edwards (1989) shows that the impact of government consumption on the real exchange rate depends on the allocation of consumption across tradables and non-tradables as well as the types of

[14] We run a simple regression of the real exchange rate on sectoral productivity differential in levels, using fixed effects estimation. We find evidence in favour of the BS hypothesis. However, the conclusion from this regression is misleading because these two variables are $I(1)$ and are not cointegrated according to Pedroni's panel cointegration tests. We then run the regression using variables in differences to avoid spurious regression. The revised results show that the sectoral productivity differential does not affect the movement of the real exchange rate.

taxes used. He assumes that government is more likely to spend more on non-tradables. Accordingly, an increase in government consumption of non-tradables will create higher demand and thus a rise in the price of non-tradables, generating a real appreciation. However, a rise in government consumption is financed by public debt and must be paid back. Hence, the government will need to increase taxes, leading to a fall in household income. This will reduce demand for non-tradables and thus cause a real depreciation. Given these two channels, the effect of an increase in government consumption is a priori indefinite and it depends on the sum of these two effects. However, in the most plausible case, the former effect is dominant and thus he expects a real appreciation in response to an increase in government consumption. On the other hand, if the assumption of the allocation of government consumption is wrong and government consumes more in tradables, an increase in government consumption will cause a persistent depreciation.

Our result shows that higher government consumption leads to a long-run real appreciation, which is in line with the most likely real exchange rate responses in the Edwards model. This suggests that government most likely allocates its consumption expenditure to non-traded goods and services, and the revenue effect is relatively small.

4.5 Forecast Error Variance Decomposition Analysis

This analysis answers the question of how much of the variance in real exchange rates over the sample period can be explained by the four shocks. Table B1 in Appendix B shows that these four shocks together can explain a moderate proportion of the forecast error variance of real exchange rates in the long run, i.e. for nine years after the shocks, the peak of the contribution of total shocks is about 23 percent. Interestingly, the trade liberalization shock can account for more than 10 percent of real exchange rate variations on impact, consistent with our expectation that trade reform will have a direct effect on real exchange rate. This also confirms results from the impulse response analysis that the trade liberalization shock has a significant effect on the real exchange rate. Thus international trade policy appears to be a powerful device for determining real exchange rate behaviour. However, trade policy might have a lower influence on real exchange rates in the long run, as the forecast error variance shares explained by this shock are smaller at longer horizons.

In addition, the forecast error variance shares explained by productivity improvement, monetary policy and government consumption shocks increase at longer horizons, suggesting that real exchange rate responses to these shocks are slow. In particular, the contribution of monetary policy shocks to the real exchange rate variance keeps increasing, and at long-term horizons such shocks contribute to fluctuations in the real exchange rate in a similar proportion as trade liberalization shocks. However, the productivity shock can account for relatively little of real exchange rate variations.

4.6 Comparison of Results with Other Approaches

We compare our results with two alternative identification strategies as below:

The pure-sign-restriction approach: This approach was developed by Uhlig (2005). Two sets of restrictions are used for this approach. First, we use the set of restrictions in Table 1 to identify the four structural shocks. Second, because Peersman (2005) mentions that the identification of other shocks should help to identify the shocks of interest, we identify not only the four shocks, but also a full set of domestic shocks. That is, in addition to extracting the four shocks of interest, we identify demand, supply and pure exchange rate shocks using an additional set of sign restrictions suggested by Farrant and Peersman (2006).

The system-penalty-function approach: This approach is a mixture of pure-sign-restriction and penalty-function approaches. It imposes more restrictions than the pure-sign-restriction approach but less restrictions than the penalty-function approach used earler in our study. In particular, instead of identifying each shock using one penalty function, the system-penalty-function approach will pick the candidate impulse matrix which minimizes the penalty function of the system for each draw of (B, Σ) from the posterior. The penalty function of the system can be written as

$$\Psi(a) = \sum_{j=1}^{4} \left[\sum_{s_j \in l_{r+}} \sum_{h=0}^{H_{re}} ff(-\frac{r_{a,s_j}(h)}{\sigma_{s_j}}) + \sum_{s_j \in l_{r-}} \sum_{h=0}^{H_{re}} ff(\frac{r_{a,s_j}(h)}{\sigma_{s_j}}) \right],$$

where j represents the j^{th} shock, and the other notation is the same as before. The penalty function of the system is the sum of the penalty functions of all four shocks. We compute the impulse responses using this approach with the set of restrictions in Table 1.

Table B2 in Appendix B summarizes the results from these experiments. It is obvious that when we use the pure-sign-restriction and the system-penalty-function approaches, most of the responses are insignificant as zero falls within the 16^{th} and 84^{th} percentile responses, excepting those for which sign restrictions are imposed. Interestingly, the responses associated with the pure-sign-restriction approach for identifying the four shocks are very similar to those for identifying a full set of shocks, in agreement with Uhlig's (2005) comment that we can concentrate on only identifying the shocks of interest.

Overall, these results give us more confidence with our identification approach. In particular, the pure-sign-restriction and the system-penalty-function approaches generate only a few significant results, probably because they relate to only a small set of the restrictions. Thus it is plausible that at any point in time, there are other existing shocks consistent with the identifying assumptions and so these two approaches induce a wide range of admissible responses. For this reason, we apply sign and zero restrictions together with four penalty functions to identify the four underlying shocks. We impose more restrictions using the penalty functions to figure out the effects of other shocks, and then achieve the responses of the shocks we require. The penalty function relies on the idea that it is likely that among all existing shocks, the shock of interest generates responses for which sign restrictions already hold. The penalty functions help us pick more decisive responses. With this approach, we successfully narrow down the range of admissible responses and we can resolve some of ambiguities in the results implied by the other approaches.

5 Conclusion

The main purpose of this paper is to investigate the short-run and long-run effects of trade liberalization, a productivity improvement in the traded sector, contractionary monetary policy and a government consumption shock on real exchange rates in eight Asian developing countries over the period 1970-2008.

We incorporate the Balassa-Samuelson effect by constructing a traded-non-traded productivity differential using a novel approach for classifying industries introduced by Dumrongrittikul (2012). We use Pedroni (1999) panel cointegration tests to examine which long-run relationships borrowed from economic theory hold among a set of our variables and then impose them in our model. Unlike previous work that has used sign-restrictions within the context of a standard structural vector autoregression, our analysis is based on a panel structural vector error correction model that is augmented with country-specific foreign variables. This allows us to account for long-run relationships among variables, and to capture the influence of foreign countries (particularly main trading partners) on a domestic country. We take the standard sign restriction approach, and also incorporate a zero restriction and a penalty-function in order to reduce the set of admissible impulse responses to a small number of similar estimates. Also we deal with the sampling uncertainty of the OLS estimates and the nonexact identification of impulse matrices by using a Bayesian method to implement the sign restrictions.

Our finding is that real GDP growth, the government consumption share and the degree of openness in the economy have a cointegrating relationship with the real exchange rate. Our impulse response analysis confirms the results expected by economic theories and found in the empirical literature so far. First, our results show that trade reform causes a significant real depreciation in the long run, and the variance decomposition analysis shows that trade liberalization is important for explaining the short-run dynamics of real exchange rates. Second, the shock on the traded-non-traded productivity differential causes an impact appreciation, although the effect of the shock on real exchange rates is short-lived. This is relevant for deliberations regarding real exchange rate behaviour among fast-growing countries in Asia, as policy makers have observed that their currencies do not tend to appreciate over the long run. Third, a contractionary monetary policy shock leads to an impact appreciation and then a depreciation back to long-run equilibrium, consistent with the Dornbusch overshooting hypothesis and long-run neutrality of monetary policy. Fourth, a rise in government consumption generates a permanent appreciation as expected.

We have seen that in Asian emerging markets with large foreign-dominated debt in particular, a dramatic depreciation of the currency can trigger a financial crisis. Policy makers should therefore monitor exchange rate fluctuations and should not pursue a policy of benign neglect on this front. Based purely on our findings, there are two suggestions for policymakers. First, policymakers need to be cautious when implementing a policy on government consumption or trade, as a change in these policies will have a permanent effect on real exchange rates. Therefore, mistimed and/or persistent effects of such a policy can make the situation worse. Second, our evidence shows that trade policy can be an effective and powerful device for dealing with real exchange rate misalignment and determining the short-run movement of real exchange rates in Asian developing countries. In contrast, a policy related to traded-sector productivity seems to be ineffective for managing real exchange rate behavior in these countries.

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A Data Appendix

For eight Asian developing countries, consumer price indices, nominal exchange rates, export value indices, import value indices, government consumption (% of GDP), exports and imports (% of GDP) were from the World Development Indicators (WDI), except export and import value indices for Thailand, India, and Pakistan that were taken from the IMF's International Financial Statistics (IFS). The IFS money market rate series were used as the short-term interest rate, except for China, for which the IFS deposit rate was used. We used GDP series measured in current and constant 1990 local currency units, classified by economic activity (ISIC 3) into seven categories. They were taken from the National Accounts Main Aggregates database, compiled by The United Nations (UN). Employment classified by ISIC 3 was from LABORSTA database. Oil prices are averages of Brent Crude series from Datastream.

B Table Appendix

Table B1: Forecast error variance shares of real exchange rates (%)

Shocks	Horizons h=0	h=2	h=4	h=6	h=8
Trade liberalization	10.55	6.74	6.34	6.46	6.69
	[56.47]	[40.02]	[36.24]	[35.42]	[35.26]
Productivity improvement	0.81	1.08	1.45	1.70	1.81
	[4.31]	[6.40]	[8.30]	[9.31]	[9.54]
Contractionary monetary policy	5.01	6.41	6.53	6.56	6.65
	[26.82]	[38.08]	[37.35]	[35.99]	[35.06]
Expansionary government consumption	2.32	2.61	3.17	3.52	3.82
	[12.40]	[15.50]	[18.11]	[19.29]	[20.13]
Total	18.68	16.83	17.49	18.23	18.96
	[100.00]	[100.00]	[100.00]	[100.00]	[100.00]

Notes: The numbers in square brackets represent the percentage of variance explained by each shock to the total variance explained by the four shocks.

Table B2: Comparison of the persistence of impulse responses produced by different approaches

Shocks	Approaches	gov	$open$	x	y	π	si	q
Trade liberalization	penalty function(1)	No	+(0-9)	-(2-9)	+(0-9)	-(0-1)	-(0-1)	+(0-9)
	penalty function(2)	No	+(0-9)	-(0-9)	+(0-9)	-(0-1)	-(0-1)	+(0-9)
	system-penalty fn.	No	+(0-9)	No	No	-(0-1)	-(1)	No
	pure sign(1)	No	+(0-9)	No	No	-(0-1)	No	No
	pure sign(2)	No	+(0-9)	-(1-9)	+(0-9)	-(0-1)	-(0-1)	+(0-1)
Productivity improvement	penalty function(1)	No	No	+(0-9)	+(0-9)	-(0-1)	-(0-1)	-(0-1)
	penalty function(2)	No	+(0-9)	+(0-9)	+(0-9)	-(0-1)	-(0-1)	No
	system-penalty fn.	No	No	+(0-9)	+(0-9)	No	No	No
	pure sign(1)	No	No	+(0-4)	+(0-9)	No	No	No
	pure sign(2)	No	No	+(0-4)	+(0-4)	No	No	No
Contractionary monetary policy	penalty function(1)	No	-(0-3)	+(1-2)	-(0-2)	-(0-8)	+(0-6)	-(0)
	penalty function(2)	No	-(0-3)	+(1-3)	-(0-2)	-(0-8)	+(0-6)	-(0)
	system-penalty fn.	No	No	+(1-3)	-(0-5)	-(0-5)	+(0-2)	No
	pure sign(1)	No	No	+(1-2)	-(0-5)	-(0-2)	+(0-3)	No
	pure sign(2)	No	No	+(1-2)	-(0-8)	-(0-1)	+(0-2)	No
Expansionary government consumption	penalty function(1)	+(0-9)	+(0-1)	-(0-9)	+(0-4)	+(0-4)	+(0-3)	-(1-9)
	penalty function(2)	+(0-9)	+(0-1)	-(0-9)	+(0-4)	+(0-4)	+(0-3)	-(1-9)
	system-penalty fn.	+(0-9)	+(0-1)	No	+(0-1)	+(0-2)	+(0-1)	-(6-9)
	pure sign(1)	+(0-9)	No	No	+(0-2)	+(0-3)	+(0)	No
	pure sign(2)	+(0-9)	+(0-1)	No	+(0-2)	+(0-4)	+(0-1)	No

Notes: 1) The penalty function refers to sign restrictions and a penalty function approach. The difference between (1) and (2) is the order of the first two shocks i.e. (1) firstly identifies a trade liberalization shock whereas (2) firstly identifies a traded-sector productivity improvement shock.

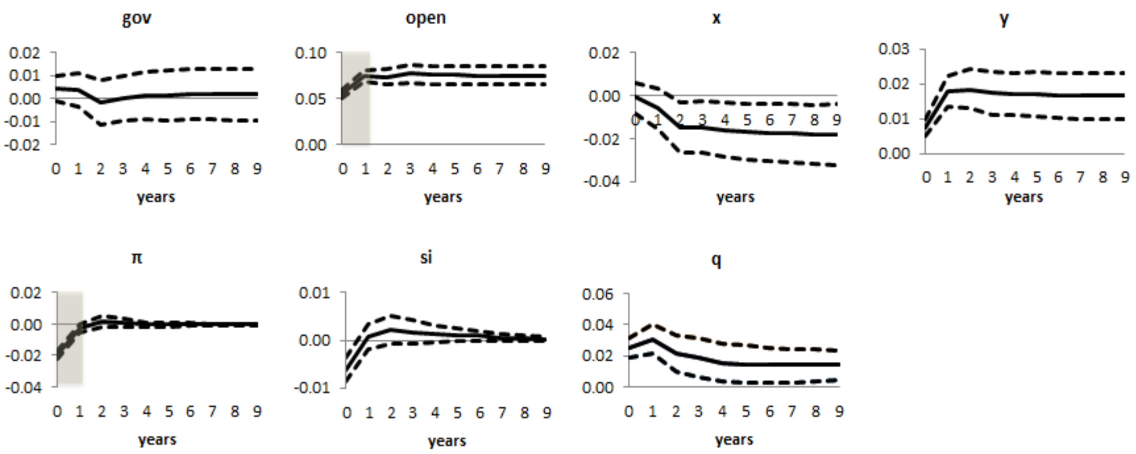
2) The pure sign represents a pure-sign-restriction approach: (1) identify four shocks, (2) identify all seven shocks in the system.

3) "No" indicates the responses that the 16-84% quantiles of the posterior distribution include zero.

4) "-" stands for a negative response and "+" stands for a positive response. The figure in parentheses is the nth period that the shock induces negative/positive responses.

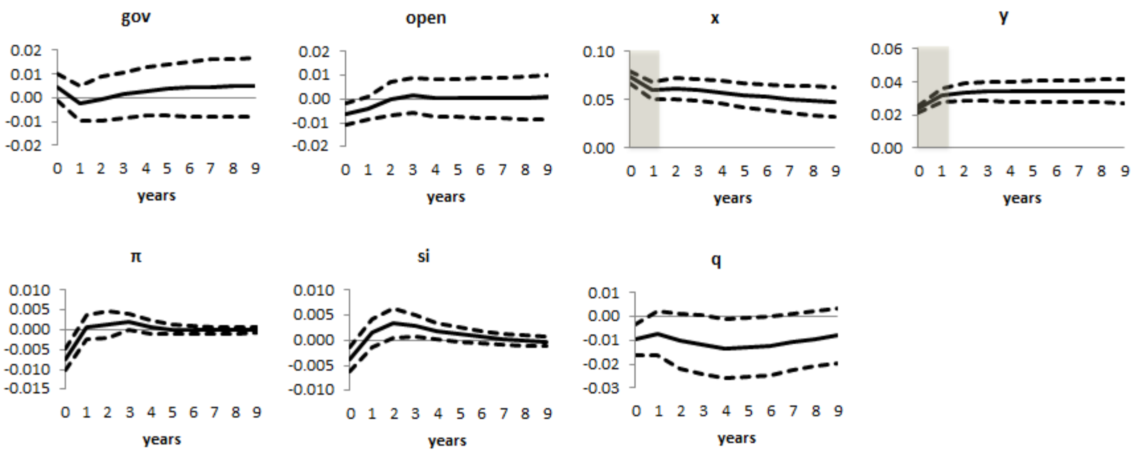
C Figure Appendix

Figure C1: Impulse responses to a trade liberalization shock of one standard deviation in size, using sign restrictions with the penalty function.



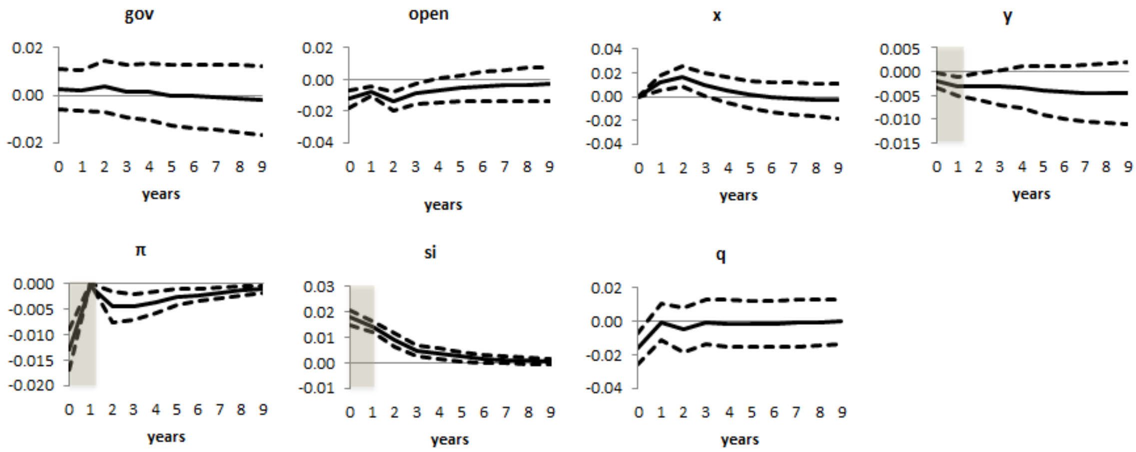
Notes: Each plot comprises of the solid line, which represents the median impulse response, and the dashed lines, which represent the 16% and 84% quantiles of the posterior distribution. The shaded areas indicate the responses restricted by sign restrictions.

Figure C2: Impulse responses to a traded-sector productivity improvement shock of one standard deviation in size, using sign restrictions with the penalty function.



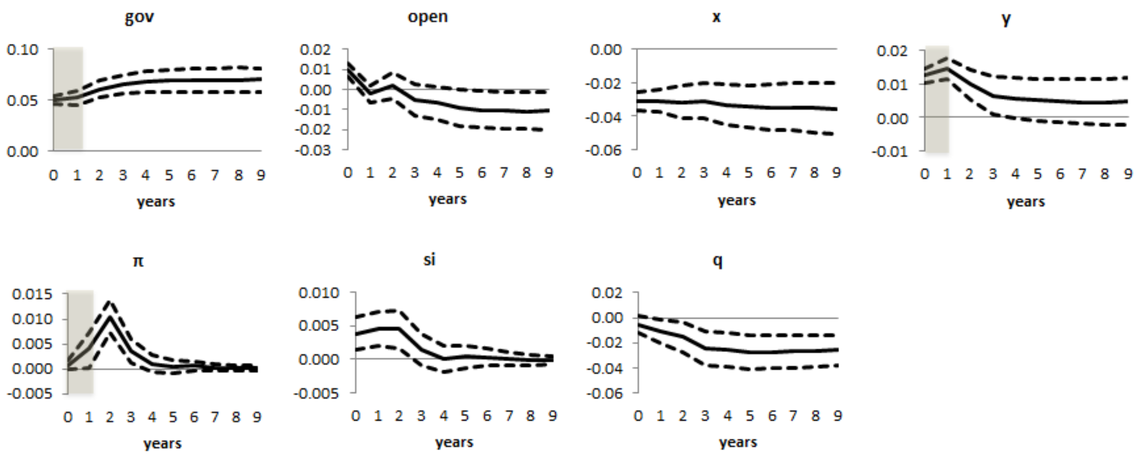
Notes: See Figure C1.

Figure C3: Impulse responses to a contractionary monetary policy shock of one standard deviation in size, using sign and zero restrictions with the penalty function.



Notes: See Figure C1.

Figure C4: Impulse responses to an expansionary government consumption shock of one standard deviation in size, using sign restrictions with the penalty function.



Notes: See Figure C1.