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 its determinants and monetary policy for eight Asian developing countries. The analysis is based on a panel Bayesian structural vector error correction model and the shocks are identified using sign and zero restrictions. We find that trade liberalization generates depreciation and higher government spending causes persistent appreciation. Traded-sector productivity gains induce appreciation; however, their effect is short-lived. Real exchange rate response to unexpected monetary tightening is consistent with the Dornbusch overshooting hypothesis and long-run neutrality of monetary policy. Strong evidence suggests that trade policy is a powerful device for driving exchange rate movements.

JEL Classification: C33, C51, E52, F31

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PREFACE

Thesis title: Real Exchange Rate Movements in Developed and Developing Economies
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Real exchange rate fluctuations may be a key factor behind the emergence of global financial crises. Recent policy discussions have emphasized the crucial role played by the real exchange rate on the economy. However, there have been small efforts to analyse empirically the fundamental sources behind real exchange rate behaviour particular in developing countries. The main purpose of this thesis is to investigate exchange rate determinants and to examine how the real exchange rate responds to shocks to its determinants, a change in policy and a global shock. Throughout our analysis, the data set includes annual time series over the period 1970-2008 and covers thirty-three countries. Asian countries are our primary interest. The following issues are addressed in the thesis:

1. Real Exchange Rate Movements and the Balassa-Samuelson Hypothesis
   The central tenet of the Balassa-Samuelson (BS) hypothesis is that high growth countries are expected to have faster productivity growth in traded sectors, so these countries will tend to experience a real appreciation. However, the BS prediction does not seem to hold in practice for fast-growing countries in Asia. This chapter aims to examine the validity of the BS hypothesis in developed and developing countries. We introduce a new approach for classifying traded and nontraded industries, use this new classification in the construction of a model that allows for the BS effect, and apply panel-data econometric methods to address this issue.

2. Do Policy-Related Shocks Affect Real Exchange Rates?
   The BS framework is extended into a general model of real exchange rates, which includes a set of theoretical real exchange rate fundamentals, policy-related variables and a common global factor. This chapter aims to examine the short-run and long-run movements of real exchange rates in response to four underlying shocks: trade liberalization, a productivity gain, monetary policy and government spending. We construct four country-group econometric models and analyse them separately. Sign restrictions with a penalty function approach are used for identifying the shocks of interest.

3. Modelling Real Exchange Rate Movements in a Global VAR Framework
   We combine four country-group models into one that is undertaken in a global vector autoregressive (GVAR) model to understand the dynamic movement of macroeconomic aggregates. The analysis places an emphasis on the reaction of real exchange rates to a sudden rise in oil prices, an unexpected monetary tightening and simultaneous shocks to productivity in the four large emerging economies in Asia, and it investigates how these shocks spill across the rest of the world. Our model is designed to account for cross-country interlinkages and it can capture many potential international transmission channels.

The thesis takes the following structure:
I: Introduction
II: Real Exchange Rate Movements in Developed and Developing Economics: A Reinterpretation of the Balassa-Samuelson Hypothesis
IV: International Linkages via Real Exchange Rates: A Global VAR Analysis
V: Conclusion

This paper is partially from the Chapter III. It focuses on the results for Asian developing countries.
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.

The main purpose of this paper is to examine the roles played by real exchange rate fundamentals and monetary policy in explaining real exchange rate behaviour, and to assess quantitatively if these factors have played an 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 responses and variance decomposition.

Our paper makes three main contributions. First, we add to the empirical literature on the relationship between the real exchange rate and its determinants for developing countries, which remains small due to the data limitations. Second, the conventional literature on sign restrictions has relied on structural vector autoregressive (SVAR) models which do not account for the long-run relationships among variables. To capture these relationships, we employ a Bayesian structural vector error correction model (SVECM) that is applied to a panel of data. We take the standard sign restriction approach, and incorporate a zero restriction and a penalty-function approach to identify four types of underlying shocks: trade liberalization, productivity improvement in the traded sector, contractionary monetary policy and expansionary government spending. Third, our sectoral productivity differential is constructed using a novel approach for classifying traded and nontraded industries introduced by Dumrongrittkul (2012). This allows us to obtain an appropriate proxy of the Balassa-Samuelson (BS) effect.

Interestingly, we find that productivity growth in the traded sector induces currency appreciation on impact as predicted by the BS model. However, its effect on real exchange rates dies out very rapidly. This is what we have seen in Asian emerging economies as their currencies do not tend to experience a real appreciation over the long run. 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
depreciation and that it also has a large contribution for real exchange rate fluctuations in the short run. This finding suggests that international trade policy should be considered as an effective and powerful instrument for dealing with real exchange rate misalignment. Second, an increase in government spending leads to persistent appreciation. Government authorities need to be aware of this when implementing a policy driven by a change in government spending 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 the long-run neutrality of monetary policy. Based purely on our finding, we suggest that in the short run, incorrectly aligned real exchange rates might potentially be corrected by focusing on a policy relating to a change in productivity, monetary policy or especially international trade policy.

The remainder of this paper is organized as follows. Section 2 briefly reviews the existing literature. Section 3 provides the theoretical background and the reasons behind choosing each real exchange rate determinant in the model. Section 4 outlines the model and the identification procedure used for recovering the shocks of interest. Section 5 presents the data set, econometric methodology and the empirical analysis. Section 6 concludes.

2 Related Literature

Early work on the relationship between the real exchange rate and its determinants has relied on cross-section comparisons or time-series techniques.\(^1\) This can lead to imprecise estimation and inconclusive hypothesis testing when data spans are short. More recent studies have turned to panel data cointegration methods. For instance, Chinn (1999) estimates a panel error correction model to examine real exchange rate behaviour in fourteen OECD countries. Chinn shows that an increase in relative traded sector productivity induces a long-run real appreciation while government spending and the terms of trade have no effects on real exchange rates. These results are inconsistent with the work of Galstyan and Lane (2009) that investigate the long-run relationship between government spending and the real exchange rate for nineteen advanced economies. Using panel Dynamic Ordinary Least Squares (DOLS) estimation, they find that government consumption induces a long-run real appreciation. Similarly, Lee, Milesi-Ferretti, and Ricci (2008) use panel DOLS to investigate real exchange rate behaviour in forty-eight industrial and emerging economies. They suggest that an increase in net foreign assets, the productivity of tradables relative to nontradables, the commodity terms of trade, the extent of trade restrictions and government consumption lead

\(^1\)Examples include Edwards (1989) and Aron et al. (1997).
to currency appreciations.

However, much of the related literature has focused on developed countries while the empirical evidence for developing countries is quite scant. Moreover, the construction of traded-nontraded productivity differentials is mostly based on arbitrary methods; that is, the classification approaches seem unreasonable because they use the same patterns of traded and nontraded sectors among industries in different countries. 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. However, the presence of a cointegrating relationship does not provide information on the direction of causality between variables.

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 Dumrongritthikul (2012) that allows for country-specific heterogeneity over each industry and changes in classification across times. 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 Long-Run Real Exchange Rate Determinants

There are several theoretical models of real exchange rate determination used for the study of developing countries, including Balassa (1964), Samuelson (1964), Edwards (1989), Obstfeld and Rogoff (1996), and Montiel (1999). Although all of these models use microfoundations with a two-sector framework (traded and nontraded sectors), some of their underlying hypotheses are different. For example, Balassa (1964), Samuelson (1964), and Edwards (1989) assume that firms in both sectors produce goods using constant returns to scale technology under perfect competition, whereas Obstfeld and Rogoff (1996) assume diminishing marginal returns to scale of a production function and monopolistic competition in the nontraded sector. Like Edwards (1989), Montiel (1999) assumes perfect competition in both sectors; however, the production function is characterized by diminishing marginal returns to scale.

Edwards (1989) shows that it is impossible to write a theoretical equation of the equilibrium real exchange rate in an explicit form without making strong assumptions that specify the exact functional forms of the revenue and expenditure; however, it can be written implicitly as a function of real exchange rate fundamentals and other exogenous variables. Given this, we do not base our empirical model on a single theoretical framework that posits that real exchange rate behaviour is a
function of some specific fundamentals, but we allow our model to include many important fundamentals suggested by various theoretical models. Specifically, we estimate a long-run relationship between the real exchange rate and a set of its fundamentals, including real GDP growth, trade liberalization, the international terms of trade, government expenditure and the sectoral productivity differential. The reduced-form equation can be written as

\[ q_{i,t} = \beta' f_{d_{i,t}} + \varepsilon_{i,t}, \]

where \( q_{i,t} \) is the log real exchange rate for country \( i \) during the period \( t \), \( f_{d_{i,t}} \) is a vector of real exchange rate fundamentals, \( \beta \) is a vector of coefficients that capture the long-run impact on \( q_{i,t} \) of changes in its fundamentals, and \( \varepsilon_{i,t} \) is an error term.

Although our set of real exchange rate fundamentals is suggested by several theoretical models, we base our analysis on the Edwards (1989) model because it is used extensively in the empirical literature of developing countries. Edwards explains the response of the equilibrium real exchange rate to changes in its fundamentals using duality theory and an intertemporal general equilibrium model of a small open economy. The economy consists of optimizing consumers and producers as well as a government with three goods: exportables, importables and nontradables. The factors of production are capital, labour and natural resources. There are two periods: the current and the future. The utility function is time separable, implying that expenditures in two periods can be substitutes. The benchmark framework of the equilibrium real exchange rate starts with the assumptions of full employment, no price rigidities, no international credit rationing, perfect foresight, perfect competition and a constant labour force; however, these assumptions can be relaxed.

Throughout our analysis, the superscript * denotes a foreign variable and no superscript implies a domestic variable. Edwards defines the real exchange rate index (\( Q \)) as the ratio of the world price of tradables (\( P_T^* \)) to the domestic price of nontradables (\( P_N \)), i.e.

\[ Q = \frac{P_T^*}{P_N}. \]

According to this definition, the real exchange rate of a small open economy depends only on the domestic price of nontradables. The real exchange rate is in equilibrium when it leads to simultaneous internal and external equilibria of the economy, given long-run equilibrium values of other relevant variables. Internal balance means that the nontradable goods market currently clears and will be in equilibrium in the future, thus implying unemployment at the natural level. External

\[ ^2 \text{However, a set of real exchange rate fundamentals is limited by data availability.} \]
equilibrium means that the current account balances in the current period and in the future will satisfy the intertemporal budget constraint i.e. the discounted sum of the current account has to be zero, and thus it is compatible with sustainable capital flow in the long run. Edwards shows that the long-term real exchange rate relies on real variables only, while both monetary and real variables can influence the short-term real exchange rate. In contrast to Purchasing Power Parity (PPP), there is not a single equilibrium value of the real exchange rate in the Edwards model, but the equilibrium real exchange rate can vary across time and it is a function of the sustainable levels of many exogenous variables.

Following mainstream literature, we define the real exchange rate as the proportion of the world price of a GDP bundle to the domestic price of the corresponding bundle:

\[
Q = NE \frac{P^*}{P} = NE \frac{(P^*_N)^\theta (P^*_T)^{1-\theta}}{(P_N)^\theta (P_T)^{1-\theta}} \propto \frac{(P^*_N)^\theta (P^*_T)^{1-\theta}}{(P_N)^\theta (P_T)^{1-\theta}},
\]

where \( NE \) is the nominal exchange rate with respect to the US, \( P \) is the general price index, \( P_T \) is the domestic price of tradables, \( P_N \) is the domestic price of nontradables, and \( \theta \) is the domestic share of nontraded goods in a GDP bundle. Due to the small open economy assumption and the law of one price for traded goods, \( P_T, P_T^* \) and \( P_N^* \) are determined abroad and they are therefore exogenous, implying that the real exchange rate in this definition depends on the same single factor as the Edwards model, i.e. the domestic price of nontradables. Therefore, although our definition of the real exchange rate is different from that of the Edwards model, this does not change the qualitative analysis of the effects of a shock to real exchange rate behaviour, and we still obtain the same conclusion as the Edwards model.

The following provides details on the reactions of equilibrium real exchange rates to a change in its fundamentals, as stated in the Edwards and BS models. Edwards analyzes the effects of a shock by using a simplified version of a general model of equilibrium real exchange rates that accounts for all essential aspects of the question we are addressing. Note that we simply provide a descriptive summary of some analysis that is associated with our study. We refer the reader to the formal mathematical and diagrammatic analysis in Edwards (1989), for more details.

3.1 Technological Progress

According to the Edwards model, the type of progress - product augmenting or factor augmenting - and the rate at which productivity improves across different sectors will have different effects on the real exchange rate. For all types of shocks to productivity gains, an income effect will generate higher demand for everything and hence an increase in the price of nontradables which,
in turn, will induce a real appreciation. In addition to this effect, technology gains will create supply effects. If the progress is of the factor augmenting type, the Rybczynski (1955) principle applies. That is, an increase in a factor of production causes more than a proportional increase in goods that are intensive in this factor, and a fall in other goods. Therefore, if a factor of production which is intensively used in the production of nontradables increases, this will lead to an increase in nontradables relative to others and a decline in their price, thus inducing a real depreciation. Moreover, if the product augmenting technological improvement increases the quantity of nontradables until there is excess supply, then the price of nontradables has to decrease to restore equilibrium. This situation will lead to a real depreciation. Therefore in the Edwards model, technological progress can cause either a real depreciation or a real appreciation depending on the domination of either supply or demand effects. Following the empirical work of Edwards, technological progress is proxied by the growth rate of real GDP in our study.

3.2 Trade Liberalization

In the Edwards model, the effect of trade liberalization such as a reduction in import tariffs is unclear 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 in both periods, 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 nontradables. This in turn entails a real depreciation. Given that the substitution effect dominates the income effect, trade liberalization will lead to a real depreciation in both periods. This is the type of reaction of the equilibrium real exchange rate that, according to Edwards and a variety of theoretical models,\(^3\) is the most likely to occur.

3.3 International Terms of Trade

Similar to other theoretical models, the reaction of the long-run real exchange rate to exogenous changes in the terms of trade in the Edwards model is a priori indefinite because the movements of the terms of trade can affect the real exchange rate through two contrary effects - income and substitution effects. That is, an improvement in the terms of trade - or an increase in the world price of exportables to importables - will induce higher real income and hence higher demand for

\(^3\)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.
nontradables. Thus this generates a higher relative price of nontradables which, in turn, leads to a real appreciation. On the other hand, under the assumption of net substitutability between importables and nontradables, a decline in the price of importables will induce higher demands for such goods, and lower the demand for nontradables. The price of nontradables then has to decrease in order to maintain equilibrium, implying a real depreciation. Therefore the long-run response of the real exchange rate to a change in the terms of trade depends on whether the intertemporal substitution or income effect dominates.

3.4 Government Expenditure

Edwards shows that the impact of government expenditure on the real exchange rate depends on the allocation of expenditure across tradables and nontradables as well as the type of taxes used. He assumes that government spends more on nontradables. Accordingly, an increase in government consumption of nontradables will create higher demand and thus a rise in the price of nontradables, generating a real appreciation. However, the excess of expenditure over revenues is financed by public debt and must be paid back. Hence, the government will need to increase taxes in either the current or future period, leading to a fall in household income. This will reduce demand for nontradables and thus cause a real depreciation. Given these two channels, the effect of an increase in government spending 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 we would expect a real appreciation in response to an increase in government spending.

3.5 Sectoral Productivity Differentials

The BS hypothesis provides an explanation of long-run real exchange rate behaviour based on productivity differentials between traded and nontraded goods. The explanation is that higher productivity in the traded sector relative to the nontraded sector leads to higher production costs for traded goods, and due to labour mobility, an increase in the wage in the traded sector raises the wage in the nontraded sector. A higher price of nontraded goods is required to maintain profitability in the nontraded sector. A higher relative price of nontraded goods to traded goods at home, relative to that of other countries then implies an appreciation of the domestic currency.

4 The Model and Identification Procedure

Throughout our analysis, all variables are in logarithmic form. The superscript * denotes a foreign variable and no superscript indicates a domestic variable. Our model consists of eight Asian devel-
oping 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. More specifically, country-specific foreign variables are constructed by using country-specific weighted averages of domestic variables of all other countries or regions. The weights are based on average trade shares over the period 2002-2008. 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. A foreign variable is given by

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

where $N (= 8)$ is the number of countries in the model, $g_{i,s,t}(g_{i,t}^*)$ is the element of $g_{i,t}(g_{i,t}^*)$ corresponding to variable $s$, and $w_{i,j}$ is the trade share of country $j$ in the total trade (imports + exports) of country $i$, such that $w_{i,i} = 0$ and $\sum_{j=1}^{26} w_{i,j} = 1$. Trade weights are used to capture the relative importance of country $j$ to country $i$, thereby, generating foreign variables which are more appropriate than traditional foreign variables, i.e. U.S. series.

### 4.1 Structural Vector Autoregression Model

We construct a model for the group of Asian developing countries, instead of an individual model for each country, so that the use of panel data methods will improve the efficiency of our parameter estimates, given the short spans of our data set. The key assumption for the purposes of estimation and inference is that foreign 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 Asian developing countries. We actually perform a formal test of the weak exogeneity of foreign variables along the lines described in Johansen (1992) and our test results support this assumption.

Based on residual serial correlation test results, we 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

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4 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.

5 Data sources for each series, and a $26 \times 26$ matrix of the trade shares used for constructing the country-specific foreign variables are available upon request.
\[ g_{i,t} = \Phi_1 g_{i,t-1} + \Phi_2 g_{i,t-2} + \Phi_3 g_{i,t-3} + \Psi_0 g^*_i+ + \Psi_1 g^*_{i,t-1} + u_{i,t}, \]

for \( t = 1, 2, ..., T \) and \( i = 1, 2, ..., N \), where \( N(=8) \) is the number of countries in the model. The notation is such that \( g_{i,t} \) is a \( k_i \times 1 \) vector of \( I(1) \) endogenous/domestic variables, \( g^*_i \) is a \( k_i^* \times 1 \) vector of \( I(1) \) weakly exogenous/country-specific foreign variables, \( \Phi_1, \Phi_2 \) and \( \Phi_3 \) are \( k_i \times k_i \) matrices of coefficients associated with lagged endogenous variables, \( \Psi_0 \) and \( \Psi_1 \) are \( k_i \times k_i^* \) matrices of coefficients associated with 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

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

where \( \Pi = (I - \Phi_1 - \Phi_2 - \Phi_3, -\Psi_0 - \Psi_1) \) 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 \( \Pi \). Suppose that the rank of \( \Pi \) is \( r_i \leq k_i \), implying that there are \( r_i \) long-run relationships among the variables. The matrix \( \Pi = \alpha \beta' \), where \( \alpha \) is a \( k_i \times r_i \) loading matrix of full column rank and \( \beta \) is a \((k_i + k_i^*) \times r_i \) matrix of cointegrating vectors of rank \( r_i \). For the estimation of the VECMX*, we will impose cointegrating restrictions \( \hat{\beta} \) computed by using panel DOLS, or \( \tilde{\beta} \) suggested by economic theory. Thus we can rewrite equation (2) as

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

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

We apply within-group estimation, i.e. we include a constant and country-specific dummy variables in the VECMX* to account for country-specific fixed effects. Alvarez and Arellano (2003) show that when \( T/N \) tends to a positive constant, the within-group estimator has negative asymptotic biases of order \( 1/T \) and these biases disappear when \( N/T \to 0 \), while the GMM estimator is asymptotically biased of order \( 1/N \). They also find that the crude GMM estimator which neglects autocorrelation in the error terms is inconsistent, despite being consistent for fixed \( T \). Moreover, Judson and Owen (1999) show that based on a root mean squared error criterion, the within-group
estimation performs better than many alternatives when \( T = 30 \). According to these findings, the within-group estimation seems to be more appropriate than other estimation techniques for our data set because the downward bias is relatively insignificant.

Given the VARX* in (1), our interest is to examine impulse responses to economically meaningful structural shocks \( v_{i,t} \). This leads to the problem of identification, given the correlation among reduced-form residuals \( u_{i,t} \). The identification issue is how to decompose reduced-form residuals \( u_{i,t} \) into structural shocks \( v_{i,t} \). 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 be of variance 1, i.e. \( E(v_{i,t}v_{i,t}') = I_{k_i} \). To obtain independence of the fundamental innovations, we find a matrix \( A \) such that \( u_{i,t} = Av_{i,t} \). We can then rewrite the reduced-form VARX* into the structural VARX* as below:

\[
A^{-1}g_{i,t} = A^{-1}\Phi_1 g_{i,t-1} + A^{-1}\Phi_2 g_{i,t-2} + A^{-1}\Phi_3 g_{i,t-3} + A^{-1}\Psi_0 g_{i,t}^* + A^{-1}\Psi_1 g_{i,t-1}^* + v_{i,t}.
\] (4)

It is clear that if we want to estimate the impulse responses of endogenous variables to structural shocks \( v_{i,t} \), we require the identifying matrix \( A \). 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 \( A \), \( a_j \).

The current set of restrictions that we have for specifying the matrix \( A \) is \( \Sigma = E[u_{i,t}u_{i,t}'] = AE[v_{i,t}v_{i,t}']A' = AA' \). Such restrictions are not sufficient to achieve a unique solution for the matrix \( A \). In particular, we need at least \( \frac{k_i(k_i-1)}{2} \) additional identifying restrictions to be imposed on the matrix \( A \).

There are three identification procedures commonly used for the orthogonalization of shocks. One procedure defines the matrix \( A \) as a lower triangular Cholesky factor of \( \Sigma \). This matrix depends on a recursive ordering of endogenous variables in the system. The second procedure imposes some structural relationships between structural shocks and reduced-form disturbances that are implied by some theoretical models or economic intuition. This can be achieved by imposing short-run and long-run restrictions e.g. restrictions on temporary and permanent components. The third procedure is a novel approach and is used in our study. This procedure identifies a shock by imposing sign restrictions on some of the impulse response functions.
4.2 Sign Restrictions with a Penalty-Function Approach

Traditional identifications are commonly based on the imposition of many stringent restrictions, most of which may not rely on theoretical considerations. To avoid this, we impose sign restrictions on the impulse response functions to identify four types of underlying disturbances; trade liberalization, productivity growth, monetary policy and government spending shocks. This new identification strategy was developed by Uhlig (2005) and extended by Mountford and Uhlig (2009). With this approach, it is not necessary to impose a priori zero contemporaneous impacts or long-run restrictions of shocks. Instead, it 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. With this approach, we also obtain results that are robust to reordering variables and selecting a particular Cholesky decomposition.

As shown by Uhlig (2005), the identification does not depend on any particular matrix $A$. On the one hand, if there exists a $k_i$-dimensional vector $m$ of unit length such that $a = \overrightarrow{Am}$, where $\overrightarrow{A} \overrightarrow{A}' = \Sigma$ and $\overrightarrow{A}$ is any arbitrary decomposition of $\Sigma$ such as a lower triangular Cholesky factor,\footnote{For the sign restriction approach, the Cholesky decomposition is only used as a computational tool, but it is not used for the purpose of identification. We can obtain similar results using any other factorization.} we can obtain an impulse vector $a$ even though the true matrix $A$ is not identified. In general, we can write $A = [a^{(1)}, \ldots, a^{(k_i)}] = \overrightarrow{A}M$ where $M = [m^{(1)}, \ldots, m^{(k_i)}]$ is an orthonormal matrix such that $MM' = I_{k_i}$. 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 $\Sigma$. This can be represented as

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

where $m_j$ is the $j^{th}$ element of $m$, and $r^{c}_j(h) \in \mathbb{R}^{k_i}$ is a $k_i \times 1$ vector of the impulse response at horizon $h$ to the $j^{th}$ shock in a Cholesky decomposition of $\Sigma$, i.e. the $j^{th}$ column of $\overrightarrow{A}$.

Moreover, it is interesting to perform variance decompositions or in other words, compute how much the 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 $s$, explained by the shock corresponding to the impulse vector $a$ can be obtained by

$$\phi_{a,s,h} = \frac{(r_{a,s}(h))^2}{\sum_{j=1}^{k_i}(r^{c}_{j,s}(h))^2},$$
where the additional index \( s \) picks the response corresponding to variable \( s \).

The sign restriction strategy can be applied to achieve the identification of \( d \) structural shocks, where \( d \leq k_i \). 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 4, rather than all impulse vectors in the matrix \( A \). In particular, we draw a \( k_i \times 4 \) matrix \( M = [m^{(1)}, m^{(2)}, m^{(3)}, m^{(4)}] \) which contains four orthonormal vectors, and then we calculate a \( k_i \times 4 \) impulse matrix as \([a^{(1)}, a^{(2)}, a^{(3)}, a^{(4)}] = \bar{A} M\). By construction, the covariances between any pair of the underlying shocks \( v_{i,t}^{(1)}, v_{i,t}^{(2)}, v_{i,t}^{(3)} \) and \( v_{i,t}^{(4)} \) corresponding to the impulse vectors \( a^{(1)}, a^{(2)}, a^{(3)} \) and \( a^{(4)} \) are zero, so we can characterize such an impulse matrix by imposing economically meaningful sign restrictions on the impulse responses.

Following Uhlig (2005), we use a Bayesian methodology 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 \((B, \Sigma)\) and a draw from a uniform distribution over the unit sphere for candidate \( m \) vectors. The Cholesky decomposition factor, \( \bar{A} \) is computed using a draw of \( \Sigma \) from the posterior. Consequently, we can calculate the candidate impulse vector as \( a = \bar{A} m \).

In this paper, 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 \((B, \Sigma)\) 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.

The penalty function suggested by Uhlig (2005) is

\[
f(ww) = \begin{cases} 
ww & \text{if } ww < 0, \\
100 \times ww & \text{if } ww \geq 0.
\end{cases}
\]  

(5)

This function is asymmetric when imposing sign restrictions, i.e. we penalize positive/wrong responses 100 times more than we reward negative/correct responses. We minimize the criterion function in order to find the best impulse vector \( a \) for each draw of \((B, \Sigma)\) from the posterior. Thus we solve

\[
a = \arg \min_{a = \bar{A} m} \Psi(a)
\]

and

\[
\Psi(a) = \sum_{l} \sum_{h=0}^{H - l} \frac{f(u_l - \frac{r_{a,l}(h)}{\sigma_l})}{\sigma_l},
\]

12
where \( l \) is the set of variables that are sign-restricted for the identification of shocks, \( r_{a,l}(h) \) is the impulse response of the \( l^{th} \) variable at horizon \( h \) to an impulse vector \( a \), and \( H_{te} \) is the last period that responses are constrained. The standard deviation of the first-differenced variable \( l \), denoted by \( \sigma_l \) 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.

Define \( \eta_l = -1 \) if \( l \) belongs to the set of variables that respond positively to a given shock according to a sign restriction, and \( \eta_l = 1 \) if \( l \) belongs to the set of variables that respond negatively to a given shock. Using numerical minimization, we can identify the first shock \( a^{(1)} = \bar{A}m^{(1)} \).

To identify the second shock, we add the restriction that the second shock is orthogonal to the first 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 \( R = [r_{1,s}^e(0), ..., r_{k,s}^e(0)] \). Therefore, additionally imposing orthogonality condition and a zero restriction, we minimize the problem below:

\[
a = \arg \min_{a=\bar{A}m,Rm=0,m^{(1)}=0} \Psi(a).
\]

### 4.3 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-nontraded 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 \( \pi \) is the inflation rate.

<table>
<thead>
<tr>
<th>Table 1: Identifying restrictions</th>
<th>( q )</th>
<th>( x )</th>
<th>( y )</th>
<th>( gov )</th>
<th>( open )</th>
<th>( si )</th>
<th>( \pi )</th>
</tr>
</thead>
<tbody>
<tr>
<td>Trade liberalization shock</td>
<td></td>
<td></td>
<td></td>
<td>+</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Productivity improvement shock</td>
<td>+</td>
<td>+</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Contractionary monetary policy shock</td>
<td>0</td>
<td></td>
<td>-</td>
<td>+</td>
<td></td>
<td>-</td>
<td></td>
</tr>
<tr>
<td>Government spending shock</td>
<td></td>
<td></td>
<td></td>
<td>+</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

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 ratio of the sum of imports and exports to GDP (\( open \)) for a year. We do not impose any restrictions on other variables as there is no theory explaining the exact responses of them to this shock.

A productivity improvement shock can be viewed as a supply shock, because it is well-known in the macroeconomic literature that a positive shock to productivity leads to a significant change in real variables. In our study, a productivity shock is identified as a shock that causes productivity in
traded sectors relative to nontraded sectors and real output to increase for a year. These restrictions correspond to the BS hypothesis which expects that productivity improvement is more rapid in countries with higher growth rates than those with lower ones.

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 a surprise increase in the nominal interest rate, there is no rise in inflation and no rise in real output over a year, whereas the sectoral productivity differential is assumed to be initially unchanged. The zero restriction on the impact response of sectoral productivity differential is plausible because productivity is normally slow to react to a nominal shock, given the sluggish behaviour of real economic activity. 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 conventional view and the standard New Keynesian model. They are also the same as those in the empirical works by Peersman (2005) and Farrant and Peersman (2006), for example.

The government spending shock is identified by restricting the responses of the government consumption to be positive for a year, while the responses of other variables are not restricted at all.

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 spending 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 it seems implausible that other shocks can have a contemporaneous effect on the open variable which is viewed as a trade-policy variable. Then we choose the shock to productivity as a second shock for identification because the productivity will take time to adjust, similar to the open variable. 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 spending shocks, we choose to order these shocks after trade liberalization and productivity shocks in order to filter out the effects of the latter shocks. This ordering also corresponds to that of Mountford and Uhlig (2009).
For computation, we find a "best impulse matrix" by undertaking the numerical minimization of the above criterion function $\Psi(a)$ on the unit sphere, given each draw of $B$ and $\Sigma$ from the posterior. We parameterize the space of unit-length vectors by using stereo projection,\footnote{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.} which is available 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 values of the total 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 $\Sigma$, we will obtain a selected impulse matrix for computing impulse responses. Then we draw a new $B$ and $\Sigma$, 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 $\Sigma$ 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, 50 and 84 percent quantiles for each of 9 step-ahead forecasts.\footnote{Note that the 16 and 84 percent quantiles give one standard deviation error bands, given that variables follow a normal distribution.}

5 Empirical Investigation

5.1 Data Description

We employ a panel data set that includes annual time series from 1970 to 2008. 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 general price index, measured by CPI, for country $i$ during the period $t$. By construction, an increase in the real exchange rate represents a real depreciation. According to the theoretical models mentioned in Section 3, the real exchange rate fundamentals are the log traded-nontraded productivity differential ($x_{i,t}$); 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.

According to Edwards (1989), nominal variables can exercise an influence on real exchange rates...
in the short run. Therefore in our empirical model, we also include the nominal short-term interest rate \( s_{i,t} = \ln(1 + NI_{i,t}/100) \) 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 \( \text{oil}_t \) 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 for all countries in our sample, all of the domestic and foreign variables in levels and all of the differences between domestic and foreign variables are approximately \( I(1) \), with the exception of nominal interest rate differentials and domestic inflation variables which are \( I(0) \).

### 5.2 Testing for the Long-Run Relationships

In the economic literature, there is a reasonable degree of consensus about the long-run relationships between variables in equilibrium; however, these relationships might or might not hold in our data set. We conduct tests for possible long-run relationships which are borrowed from economic theory. Based on the long-run properties of macroeconomic models, we consider the following well-known long-run relationships as possible candidates:

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

**Combination of the Balassa-Samuelson and Edwards models**

\[
q_{i,t} - \lambda_1(x_{i,t}^* - x_{i,t}) - \lambda_2(y_{i,t}^* - y_{i,t}) - \lambda_3(gov_{i,t}^* - gov_{i,t}) - \lambda_4(tt_{i,t}) - \lambda_5(open_{i,t}) \sim I(0)
\] (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 \( s_{i,t} \sim I(0) \). The third relationship represents the relative output convergence condition derived from the Solow-Swan neoclassical growth model. 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 \( si_{i,t} - si^*_{i,t} \sim I(0) \). We focus on the fifth relationship i.e. the long-run relationship between the real exchange rate and its fundamentals. 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). As mentioned before, \( q_{i,t} \) and \( si_{i,t} \) are approximately \( I(1) \), suggesting that PPP and the Fisher Equation do not hold in all countries.

In order to test for the long-run relationships (3)-(5), we apply four residual-based cointegration tests suggested by Pedroni (1999).\(^{10}\) 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.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Panel PP</th>
<th>Group PP</th>
<th>Panel ADF</th>
<th>Group ADF</th>
</tr>
</thead>
<tbody>
<tr>
<td>( y_{i,t}, y^*_{i,t} )</td>
<td>0.32</td>
<td>0.48</td>
<td>0.03</td>
<td>0.03</td>
</tr>
<tr>
<td>( si_{i,t}, si^*_{i,t} )</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>( q_{i,t}, (x^<em><em>{i,t} - x</em>{i,t}), (y^</em><em>{i,t} - y</em>{i,t}), (gov^*<em>{i,t} - gov</em>{i,t}), t_i, open_{i,t} )</td>
<td>0.86</td>
<td>0.94</td>
<td>0.88</td>
<td>0.85</td>
</tr>
</tbody>
</table>

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 the UIP condition. 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.

There is not enough evidence in favour of 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, in order to ascertain that there is no cointegrating relationship (5), we conduct a parsimonious approach. That is, we test all possible

\(^{10}\) 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.
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 3 for the test results.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Panel PP</th>
<th>Group PP</th>
<th>Panel ADF</th>
<th>Group ADF</th>
</tr>
</thead>
<tbody>
<tr>
<td>(q_{i,t}, (y_{i,t}^* - y_{i,t}), (gov_{i,t}^* - gov_{i,t}), open_{i,t})</td>
<td>0.08</td>
<td>0.09</td>
<td>0.01</td>
<td>0.00</td>
</tr>
</tbody>
</table>

Notes: See Table 2.

5.3 Estimating and Interpreting the Cointegrating Vector

We use panel DOLS to estimate the cointegrating vector, because the DOLS estimator has smaller size distortions and outperforms the OLS and the fully modified OLS (FMOLS) estimators in both finite and infinite samples, according to the findings of Kao and Chiang (1999).\(^{11}\)

As suggested by Mark and Sul (2003), we apply within-dimension DOLS estimation. 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, and allowing for a limited degree of cross-sectional dependence through time-specific effects seems sufficient to capture the heterogeneity in the country members of the panel. With the within-dimension technique, 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 DOLS regression for estimating a cointegrating vector of the real exchange rate relationship is given by

\[ q_{i,t} = d_i + \theta_t + \beta' f_{d_{i,t}} + \delta_i s_{d_{i,t}} + \varepsilon_{i,t}, \]

where \(i = 1, 2, ..., N; f_{d_{i,t}} = ((y_{i,t}^* - y_{i,t}), (gov_{i,t}^* - gov_{i,t}), open_{i,t})'; s_{d_{i,t}} = (\Delta f_{d_{i,t-1}}, \Delta f_{d_{i,t}}, \Delta f_{d_{i,t+1}})'; d_i\) is a country-specific effect; \(\theta_t\) is a common time-specific factor which is used to capture some forms of cross-sectional dependence across countries, as the asymptotic distribution theory requires the cross-sectional independence of residuals. We focus on the estimate of \((1, -\beta')'\) which is a cointegrating vector between \(q_{i,t}\) and \(f_{d_{i,t}}\) and is allowed to be identical across individual countries in the panel. The coefficients \(\beta'\) capture the long-run impact on \(q_{i,t}\) of long-run changes in the \(f_{d_{i,t}}\) variables.

\(^{11}\)It is well-known that the OLS estimator is typically biased in panel data settings, but although the FMOLS estimator was created to solve the bias of OLS estimators by using a complicated estimation technique to account for the endogeneity and serial correlation, Kao and Chiang (1999) show that it is still biased in finite samples and does not improve over the OLS estimator in general cases.
Table 4: The panel DOLS estimates of the real exchange rate relationship

<table>
<thead>
<tr>
<th>Variables</th>
<th>$y_{i,t} - y_{i,t}$</th>
<th>$gov_{i,t} - gov_{i,t}$</th>
<th>open_{i,t}</th>
</tr>
</thead>
<tbody>
<tr>
<td>Estimated coefficients</td>
<td>$-0.539^{**}$</td>
<td>0.276**</td>
<td>0.071</td>
</tr>
<tr>
<td>$F$-statistic</td>
<td>15.064**</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: ** indicate 1% significance levels. Standard errors are reported in parentheses.

Table 4 shows that this set of real exchange rate fundamentals influences real exchange rate behaviour, as the $F$-statistic is significantly different from zero. Also 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 spending 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.

We note that although the estimates of the cointegrating vector can represent the relationship between the real exchange rate and its fundamentals, these estimates do not imply what the direction of causation is. Therefore our next task is to examine the direction of causation, and to determine how quickly shocks to these fundamentals affect real exchange rates and other variables, how large these impacts are, and how long these variables, especially real exchange rates, take to return to their long-run equilibrium.

5.4 Impulse Response Analysis

The VECM$x^*$ comprises seven endogenous variables and six weakly exogenous variables as follows: $^{12}$

$$g_{i,t} = [q_{i,t}, x_{i,t}, y_{i,t}, gov_{i,t}, open_{i,t}, s_{i,t}, \pi_{i,t}]' \text{ and }$$

$$g^*_{i,t} = [x^*_{i,t}, y^*_{i,t}, gov^*_{i,t}, s_{i,t}, \pi^*_{i,t}, oil_t]' .$$

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 VECM$x^*$.$^{13}$

We mainly focus on the responses of the real exchange rate to four structural innovations - trade liberalization, productivity growth, contractionary monetary policy and government spending

---

$^{12}$ Note that we have dropped the terms of trade ($t_{i,t}$) 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.

$^{13}$ 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.
shocks. Figure 1 shows the median impulse responses of the real exchange rate together with the 16th and 84th percentile responses over nine years after the shocks. Shocks are normalized to a magnitude of one standard deviation in size. 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.](image)

Fry and Pagan (2011) point out that although the median of the impulse responses is popular in the sign-restriction literature, it may not be a good way of summarizing information. Therefore, we also use a different approach suggested by Fry and Pagan (2011) and find a single model whose impulse responses are as close to the median responses as possible. However, we find that these two different approaches do not produce any significantly different results (See Figure 1). 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.\(^\text{14}\)

**5.4.1 Trade Liberalization Shock**

According to the Edwards model, when a small country permanently liberalizes foreign trade or lowers trade barriers, demand for importables increases. On the other hand, this entails a decline in demand for nontradables as the relative price for nontradables rises. Assuming the Marshall-Lerner condition holds, a real depreciation is required to restore internal and external equilibrium. As shown in Figure 1, our finding provides evidence in support of this view. That is, controlling for other factors - such as sectoral productivity differentials, economic growth, government spending shares, and capital flows - a permanent trade liberalization results in a real depreciation of the domestic currencies which persists for about three years.

\(^{14}\)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, except that real exchange rate responses to a government spending shock are stronger.
Moreover, Figure C1 shows that the shock causes a rise of real output, probably because a trade reform provides incentives for domestic firms to compete or seek new markets overseas, encouraging them to invest more. This also leads to an increase in nominal interest rates. Further, a positive income effect caused by the shock results in higher demand for all goods and raises their prices.

5.4.2 Productivity Improvement Shock

According to the BS hypothesis, if productivity in the traded sector in one country grows faster than that in other countries (given that productivity growth in the nontraded sector for all countries is slow), its currency will persistently appreciate. However, our result shows that the productivity shock causes a contemporaneous appreciation and its effect dies out in the short run. This is what we have actually seen in rapidly growing countries in Asia such as China and India, since they do not seem to experience a long-run appreciation.

Our results differ from the traditional BS model and the related empirical literature because of two main reasons. First, according to the BS hypothesis, the traded-nontraded 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. Second, 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 nontraded 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. In our study, we use a novel classification approach introduced by Dumrongrittitkul (2012) in which classifications change across countries and periods. Nevertheless, we still find that a productivity differential shock only causes an impact appreciation.\textsuperscript{15}

Otherwise, our results are in line with traditional trade theory, as it is known that productivity gains in the traded sector can enhance the capacity of domestic firms to compete in international markets, leading to a rise in international trade share after the shock.

\textsuperscript{15} 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.
5.4.3 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. Following directly from these, the Dornbusch overshooting hypothesis suggests that ceteris paribus an unanticipated tightening in monetary policy leads to an impact appreciation beyond its long-run value followed by a depreciation toward the terminal value, in line with UIP.

In our study, the price puzzle and the output puzzle do not occur, 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 the long-run neutrality of monetary policy (See Figure C3).

5.4.4 Government Spending Shock

Following the traditional view, an increase in government consumption is regarded as a rise in the relative demand for nontraded goods, thereby increasing their prices and leading to a real appreciation. Our result shows that higher government spending leads to a long-run real appreciation, which is in line with the prediction of the theoretical models like the Edwards model and the existing literature e.g. Chinn (1999), Galstyan and Lane (2009).

In the literature, some researcher consider the government spending shock as a fiscal policy shock since a higher government consumption share can indicate expansionary fiscal policy. However, fiscal policy actually relies on two main components: government spending and government revenue. Therefore, the use of government spending as a proxy of fiscal policy can lead to misleading conclusions. This can be easily seen from our results, i.e. a higher government spending share causes a fall in interest rates and inflation. This may be because an increase in the government spending share in these countries is a result of contractionary fiscal policy, not expansionary fiscal policy, or in other words, a rise in the government consumption share may be due to higher government revenue through an increase in tax. Accordingly, a fall in private consumption (negative
income effect) due to higher tax may cause deflation and a drop of output as shown in Figure C4.

5.5 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 sizable proportion of the forecast error variance of real exchange rates, i.e. for nine years after the shocks, the peak of the contribution of total shocks is about 32 percent. Interestingly, the trade liberalization shock can account for more than 20 percent of real exchange rate variations on impact, confirming results from the impulse response analysis that the trade liberalization shock has a significant effect on the real exchange rate. This provides evidence that international trade policy can indeed be a powerful device for determining real exchange rate behaviour in the short run. 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 and monetary policy shocks increase at longer horizons, suggesting that real exchange rate responses to these two shocks are very slow. In particular, the contribution of monetary policy shocks to the real exchange rate variance keeps increasing, and at long-term horizons such a shock contributes the most to fluctuations in the real exchange rate.

5.6 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 as before. 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 of interest, but also a full set of 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 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 \((B, \Sigma)\) from the posterior. Therefore, it is not necessary that for each shock, the largest responses of the variables for which sign restrictions are imposed, are obtained.

The penalty function of the system can be written as

\[
\Psi(a) = \sum_{j=1}^{4} \sum_{t} \sum_{h=0}^{H_{zz}} f(u) \frac{r_{a,j}(h)}{\sigma_t},
\]

where \(j\) represents the \(j^{th}\) shock, and the other notation is the same as before. It is clear that the penalty function of the system is the sum of the penalty functions of all the 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, almost all 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 there is no reason to identify a full set of shocks, or in other word, 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. This can be viewed as an illustration of an identification problem due to weak information, 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.
6 Conclusion

The main purpose of this paper is to explore the role of a set of exchange rate fundamentals and an unexpected change in monetary policy on real exchange rate behaviour in eight Asian developing countries over the period 1970-2008.

We incorporate the Balassa-Samuelson effect by constructing a traded-nontraded productivity differential using a novel approach for classifying industries, introduced by Dumrongrittikul (2012). The long-run relationship between the real exchange rate and its fundamentals is examined by using the Pedroni (1999) panel cointegration tests and panel dynamic ordinary least squares estimation suggested by Mark and Sul (2003). The most innovative aspect of our study is that we take the standard sign restriction approach based on a structural vector autoregression model, and incorporate a zero restriction and a penalty-function approach as well. We apply this identification method to a Bayesian structural vector error correction model of a panel to extract four underlying structural shocks - trade liberalization, a productivity improvement in the traded sector, contractionary monetary policy and a government spending shock. This approach is superior to other identification methods because it requires only a small set of a priori theoretical restrictions and it generates results that are robust to the choice of reordering variables and selecting a particular variance decomposition. It is also robust to non-stationarity of the series, including structural breaks and regime changes.

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. Firstly, our results show that a permanent trade reform causes a significant real depreciation, and the variance decomposition analysis shows that trade liberalization is very important for explaining the short-run dynamics of real exchange rates. Secondly, the shock on the traded-nontraded productivity differential causes an impact appreciation; however, the effect of the shock on real exchange rates is short-lived over the short run. 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. Thirdly, 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 the long-run neutrality of monetary policy. Fourthly, a rise in government spending generates a real appreciation in the long run.

We have seen that in Asian emerging markets with large foreign-dominated debt in particular, a
dramatic depreciation of the currency could 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 an appropriate policy response to potential misalignments of real exchange rates. First, a policy driven by government spending and a monetary policy can affect both the real economy and financial markets, so government and monetary authorities need to be cautious in implementing these policies to constrain real exchange rate misalignment. In particular, a change in government spending will have a long-run effect on real exchange rates, so mistimed and/or persistent effects of the policy can make the situation worse. Second, our evidence shows that shocks to productivity improvement in traded sectors, monetary policy and trade liberalization principally have a contemporaneous impact on real exchange rates. Therefore we suggest that in the short run, incorrectly aligned real exchange rates might potentially be corrected by focusing on a policy relating to a change in productivity, monetary policy or especially international trade policy. Strong evidence shows that trade policy can be an effective and powerful device for dealing with real exchange rate misalignment and determining the short-run dynamics of real exchange rates in Asian developing countries.

References


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 Bank, 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). Oil prices are averages of Brent Crude series from Datastream.
B Table Appendix

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

<table>
<thead>
<tr>
<th>Shocks</th>
<th>Horizons</th>
<th>h=0</th>
<th>h=2</th>
<th>h=4</th>
<th>h=6</th>
<th>h=8</th>
</tr>
</thead>
<tbody>
<tr>
<td>Trade liberalization</td>
<td></td>
<td>20.03</td>
<td>8.73</td>
<td>6.01</td>
<td>5.13</td>
<td>4.66</td>
</tr>
<tr>
<td></td>
<td></td>
<td>[72.95]</td>
<td>[49.63]</td>
<td>[38.12]</td>
<td>[33.26]</td>
<td>[30.21]</td>
</tr>
<tr>
<td>Productivity improvement</td>
<td></td>
<td>0.72</td>
<td>1.13</td>
<td>1.53</td>
<td>1.77</td>
<td>1.89</td>
</tr>
<tr>
<td></td>
<td></td>
<td>[2.63 ]</td>
<td>[6.41 ]</td>
<td>[9.68 ]</td>
<td>[11.47 ]</td>
<td>[12.24]</td>
</tr>
<tr>
<td>Contractionary monetary policy</td>
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<td>4.36</td>
<td>6.10</td>
<td>6.63</td>
<td>6.86</td>
<td>7.08</td>
</tr>
<tr>
<td></td>
<td></td>
<td>[15.88]</td>
<td>[34.68]</td>
<td>[42.06]</td>
<td>[44.52]</td>
<td>[45.88]</td>
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<tr>
<td>Government spending</td>
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<td>2.34</td>
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<td>1.60</td>
<td>1.66</td>
<td>1.80</td>
</tr>
<tr>
<td></td>
<td></td>
<td>[8.53 ]</td>
<td>[9.29 ]</td>
<td>[10.14]</td>
<td>[10.76]</td>
<td>[11.67]</td>
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<tr>
<td>Total</td>
<td></td>
<td>27.45</td>
<td>17.58</td>
<td>15.76</td>
<td>15.42</td>
<td>15.43</td>
</tr>
<tr>
<td></td>
<td></td>
<td>[100.00]</td>
<td>[100.00]</td>
<td>[100.00]</td>
<td>[100.00]</td>
<td>[100.00]</td>
</tr>
</tbody>
</table>

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: Comparision of the persistence of impulse responses produced by different approaches

<table>
<thead>
<tr>
<th>Shocks</th>
<th>Approaches</th>
<th>gov</th>
<th>open</th>
<th>x</th>
<th>y</th>
<th>π</th>
<th>si</th>
<th>q</th>
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</thead>
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<tr>
<td>Trade liberalization</td>
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<td>+0(9)</td>
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<td>+0(2)</td>
<td>+0(2)</td>
<td>+0(2)</td>
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<tr>
<td></td>
<td>penalty function(2)</td>
<td>-0(1)</td>
<td>+0(9)</td>
<td>No</td>
<td>+0(1)</td>
<td>+0(2)</td>
<td>+0(2)</td>
<td>+0(2)</td>
</tr>
<tr>
<td></td>
<td>system-penalty fn.</td>
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<td>+0(9)</td>
<td>No</td>
<td>No</td>
<td>+1(1)</td>
<td>+1(1)</td>
<td>No</td>
</tr>
<tr>
<td></td>
<td>pure sign(1)</td>
<td>No</td>
<td>+0(9)</td>
<td>No</td>
<td>No</td>
<td>No</td>
<td>No</td>
<td>No</td>
</tr>
<tr>
<td></td>
<td>pure sign(2)</td>
<td>No</td>
<td>+0(9)</td>
<td>No</td>
<td>No</td>
<td>No</td>
<td>No</td>
<td>No</td>
</tr>
<tr>
<td>Productivity improvement</td>
<td>penalty function(1)</td>
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<td>+(1-4)</td>
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<td>+0(9)</td>
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<td>-0(1)</td>
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<tr>
<td></td>
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<td>No</td>
<td>+(1-9)</td>
<td>+0(9)</td>
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<td>-0(1)</td>
<td>-0(1)</td>
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<tr>
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<td>No</td>
<td>+0(9)</td>
<td>+0(9)</td>
<td>-0(1)</td>
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<tr>
<td></td>
<td>pure sign(1)</td>
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<td>No</td>
<td>+0(9)</td>
<td>+0(9)</td>
<td>No</td>
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<td>No</td>
</tr>
<tr>
<td></td>
<td>pure sign(2)</td>
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<td>No</td>
<td>+(3-9)</td>
<td>+(3-9)</td>
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<tr>
<td>Contractionary monetary policy</td>
<td>penalty function(1)</td>
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<td>No</td>
<td>+(1-2)</td>
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<td>-0(8)</td>
<td>+0(6)</td>
<td>-0(0)</td>
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<tr>
<td></td>
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<td>No</td>
<td>No</td>
<td>+(1-2)</td>
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<td>-0(8)</td>
<td>+0(6)</td>
<td>-0(0)</td>
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<tr>
<td></td>
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<td>-0(5)</td>
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<tr>
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<td>No</td>
<td>No</td>
<td>-0(9)</td>
<td>-0(4)</td>
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<td>Government spending</td>
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<td>No</td>
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<td>-0(1)</td>
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<td>-0(9)</td>
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<tr>
<td></td>
<td>pure sign(2)</td>
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<td>No</td>
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</table>

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 n th period that the shock induces negative/positive responses.
5) Bold letter indicates the response restricted by sign restrictions.
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 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.

![Graphs showing impulse responses for various variables over time.](image)

Notes: See Figure C1.

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

![Graphs showing impulse responses for various variables over time.](image)

Notes: See Figure C1.