

How Do Shocks to Domestic Factors 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 for fourteen Asian developing countries. The analysis is based on a panel 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 not immediate and last only for a few years. Real exchange rate responses to unexpected monetary tightening are consistent with the long-run neutrality of money. The evidence suggests that trade liberalization and government consumption have a strong effect on real exchange rates, while the effects of traded-sector productivity shocks are much weaker.

Keywords: Exchange rate fundamentals, Government consumption, Monetary policy, Panel vector error correction model, Productivity improvement in the traded sector, Real exchange rates, Sign and zero restrictions, Trade liberalization.

JEL classification: C33, C51, E52, F31

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1 Introduction

Emerging Asia's participation in global economic activity has received increasing attention in recent times, with offshore production and the outsourcing of services playing important roles in the rapid industrialization of countries such as China and India, and markets in many developing Asian countries now becoming highly competitive from an international perspective. Supportive domestic policies have opened up many Asian economies, stimulating international financial flows and associated investment, and allowing trade links between Asia and the rest of the world to become strong. Indeed, the International Monetary Fund (IMF, 2007) reported that emerging Asia's share of global trade flows reached more than a third (i.e. 34%) in 2006. Asian countries have also become increasingly integrated via intraregional trade that reflects a rise in inter-industry processing and vertically integrated production (Gupta 2012), with regional trade flows averaging annual increases of over 10 percent (in value-added terms) since 1990 (IMF, 2014).

The increasing influence of Asia in the world economy coupled with greater regional economic integration between Asian countries points to the growing importance of Asian real exchange rates and associated policies. A stable real exchange rate is key to successful outward-oriented, export-based development strategies, and exchange rates that are poorly aligned with fundamentals can lead to widespread macroeconomic and financial instability in developing countries, as discussed extensively in Edward's (1988a) book. Thus, ongoing empirical research on what constitutes a misalignment of exchange rates from long-run equilibrium and how macroeconomic policies can influence real exchange rate movements is important, especially for the rapidly changing and growing economies within Asia. Further, while several Asian countries such as China, India and Vietnam are under political pressure to raise the value of their currencies relative to the USD, it is appropriate to ask how this might be best achieved, should these countries decide to pursue this objective.¹

Economists suspect that macroeconomic fundamentals such as productivity, trade liberalization, government consumption and interest rates play important roles in the determination of Asian real exchange rates, just as they do for developed economy exchange rates, but the relative importance of these fundamentals and the mechanisms via which they operate are likely to differ in developed and developing economy environments, and indeed differ in different parts of the world. There are a few empirical studies that specifically focus on Asian exchange rates (some examples include Chinn (2000) and Drine and Rault (2004)), but this literature is sparse, and does not account for more recent trends in interaction between developing Asian countries and the rest of the world.

¹See Rodrick (2008) and Schröder (2013) for discussion on the issue of whether undervaluation or overvaluation of developing country real exchange rates promote or hinder growth, and see Morrison and Labonte (2013) for detailed discussion on the ramifications of revaluing the Renminbi for the United States and China.

This paper aims to provide a contemporary examination of the roles played by trade liberalization, productivity improvement in the traded sector, contractionary monetary policy and expansionary government consumption in explaining real exchange rate movements in developing Asian countries, and to assess which of these factors (if any) have played an important driving role. Our broad aim is to provide an empirical framework that can inform policy debates on Asian exchange rate movements. Our primary data set comprises annual macroeconomic and financial variables for fourteen developing Asian countries over the period from 1970 to 2008, and we supplement this with aggregated data from the rest of the world. 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 in order to trace out both short and longer run patterns associated with exchange rate responses to changes in fundamentals.

Our paper makes four main contributions. First, we add to and update the empirical literature on the relationship between the real exchange rate and its determinants for developing Asian countries, which has been limited up until now by the lack of sufficient data. Second, we depart from the conventional empirical literature on shock identification in open economy models by using an augmented structural vector error correction model (SVECMX*) that we apply to a panel of data. This model not only accounts for long-run relationships between Asian country specific variables as well as the short-run relationships that are typically modelled within a standard structural vector autoregression (SVAR), but it also allows us to capture the aggregated influence of foreign country variables. Third, we impose sign restrictions and a zero restriction on impulse responses, using a penalty function that picks out a small set of admissible impulse responses, and a Bayesian approach that allows for parameter estimation uncertainty and deals with nonexact identification of impulse vectors. Our use of sign restrictions allows us to define economically meaningful policy shocks. Fourth, our sectoral productivity differential is constructed using a new classification of traded and nontraded industries proposed by Dumrongrittikul (2012). This classification uses trade and price data to produce time-varying country specific classifications of industries, and we use it to construct a measure of productivity in the traded sector relative to the nontraded sector. We use this new productivity variable to assess the Balassa (1964) - Samuelson (1964) (BS) hypothesis.

We find that productivity improvements in the traded sector induces a real appreciation as predicted by the BS hypothesis, but that this effect on real exchange rates is not immediate and it dies away within five years. 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 (1988b, 1989) and other theoretical models. First, there is strong evidence that trade liberalization generates permanent depreciation and that it also makes a large contribution to real exchange rate fluctuations, especially in the short run. This suggests that most of the variation in real exchange rates is determined by trade. Second, increased government consumption leads to persistent appreciation. Governments need to be aware of this when using fiscal policy to influence the economy. Third, we find that money policy has no long-run effect on real exchange rates.

The remainder of this paper is organized as follows. Section 2 briefly reviews the existing empirical literature. Section 3 provides the theoretical background. 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

There are two main strands of related literature. The first strand is the work on the determinants of real exchange rates. Early empirical work on this topic has relied on cross-section comparisons or standard time-series techniques. This can lead to imprecise estimation and inconclusive hypothesis tests when samples are small. More recent studies have turned to panel data cointegration methods. 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, Ricci et al. (2013) use panel dynamic ordinary least squares (DOLS) estimation and focus on forty-eight industrial and emerging economies. Their work suggests that increases in net foreign assets, the productivity of tradables relative to nontradables, the commodity terms of trade, trade restrictions and government consumption cause currency to appreciate.

The second strand of literature which is closely related to our study is the empirical work on theoretically-motivated sign restrictions on impulse responses. Sign restrictions have become increasingly popular in recent years because they avoid the use of strong identifying restrictions. They have been used to examine monetary policy shocks (e.g. Faust and Rogers (2003), Farrant and Peersman (2006), and Scholl and Uhlig (2008)), to measure the effects of government spending shocks (e.g. Mountford and Uhlig (2009) and Pappa (2009)), and to study productivity shocks (e.g. Dedola and Neri (2007), Corsetti et al. (2014), and Peersman and Straub (2009)).

Most previous literature on exchange rate determination has focused on developed countries and does not aim to explain the effects of shocks on real exchange rates. In addition, the existing papers

typically examine the causal effects of exchange rate determinants on real exchange rates by relying only on the study of cointegrating relationships. Shortcomings associated with this approach are that the presence of a cointegrating relationship does not provide information on the direction of causality, and study of cointegration does not cast light on short-run dynamics.

In this paper, we focus on real exchange rate behavior in developing Asian countries. The 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 the new classification introduced by Dumrongritikul (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 Theoretical Considerations

Several theoretical models have been introduced to examine real exchange rate behavior of developing countries in responses to changes in its determinants. Examples include Balassa (1964), Samuelson (1964), Dornbusch (1976), Edwards (1988b, 1989), and Obstfeld and Rogoff (1996). Although all of these models use microfoundations with a two-sector framework (traded and non-traded sectors), some of their underlying hypotheses are different. The following sub-sections briefly discuss models that are widely used to examine the responses of the real exchange rate to the shocks of interest, i.e. trade liberalization, government consumption, productivity and monetary policy.

3.1 Real exchange rates in a small open economy

Edwards (1988b) develops a dynamic model for a small open developing economy that has a dual nominal exchange rate system, and he uses it to investigate the effects of various policies (including trade controls) that can influence exchange rates. This model incorporates a freely determined exchange rate for financial transactions as well as a controlled exchange rate for commercial transactions,² and the latter allows for real exchange rate misalignment and a distinction between long-run and short-run movements in the real exchange.

The model assumes that the economy consists of consumers, producers and a government, it produces exportables and nontradables, and it consumes importables and nontradables. The dual exchange rate system includes a fixed nominal exchange rate (E) for commercial transactions and a freely floating nominal exchange rate (δ) for financial transactions, which give rise to an exchange

²The freely determined exchange captures "black markets" that are often present in developing countries.

rate spread of $\rho = \delta/E$. Total assets (A) of this country in domestic currency is the sum of domestic money (M) and foreign money (δF), and although there is no international capital mobility ($\dot{F} = 0$) there is a positive initial stock of foreign money (F_0). A key assumption is that there is a negative relationship between the desired ratio of real domestic money ($m = M/E$) to real foreign money (ρF) and the rate of depreciation of the free rate ($\dot{\delta}/\delta$); i.e. $\frac{m}{\rho F} = \sigma(\frac{\dot{\delta}}{\delta})$ where $\frac{\partial \sigma}{\partial \frac{\dot{\delta}}{\delta}} < 0$.

The model incorporates an import tariff (τ), with its revenues being redistributed back to the public in a lump-sum fashion. Normalizing the price of exportables in terms of foreign currency to be one ($P_X^* = 1$), letting P_N be the price of nontradables and P_M (P_M^*) be the price of importables that includes (excludes) the import tariff, we have $P_M = EP_M^* + \tau$, and for a government that consumes G that consists of importables (G_M) and nontradables (G_N), we have $G = P_N G_N + EP_M^* G_M$. Real government consumption (g) can be written in terms of exportables as

$$g = g_N + g_M,$$

where $g = \frac{G}{E}$, $g_N = \frac{G_N P_N}{E}$ and $g_M = P_M^* G_M$. Government spending is financed via nondistortionary taxes (t) so that a stationary equilibrium implies that $G = t$. If domestic credit creation (i.e. $\dot{D} > 0$) is sustainable given the rate of change in commercial exchange (\dot{E}/E), then $G = t + \dot{D}$.

Private demand for importables (C_M) and nontradables (C_N) are modelled as functions of the relative price of importables to nontraded goods (i.e. $e_M = P_M/P_N$) and the level of real assets in terms of exportables (i.e. $a = \frac{A}{E}$). Thus

$$\begin{aligned} C_M &= C_M(e_M, a), \text{ where } \frac{\partial C_M}{\partial e_M} < 0, \frac{\partial C_M}{\partial a} > 0; \text{ and} \\ C_N &= C_N(e_M, a), \text{ where } \frac{\partial C_N}{\partial e_M} > 0, \frac{\partial C_N}{\partial a} < 0. \end{aligned}$$

The supply for both goods (Q_X and Q_N) is a function of $e_X = E/P_N$, which measures the price of exportables ($P_X^* = 1$) relative to nontraded goods. This gives rise to

$$\begin{aligned} Q_X &= Q_X(e_X), \text{ where } \frac{\partial Q_X}{\partial e_X} > 0; \text{ and} \\ Q_N &= Q_N(e_X), \text{ where } \frac{\partial Q_N}{\partial e_X} < 0. \end{aligned}$$

The model defines the current account in terms of foreign currency as

$$CA = Q_x(e_X) - P_M^* C_M(e_M, a) - P_M^* G_M,$$

and the stock of international reserves in foreign currency (R) accumulates according to $\dot{R} = CA$.

The domestic stock of money then accumulates according to $\dot{M} = \dot{D} + E\dot{R}$. Finally, the model defines the real exchange rate as

$$q = \frac{E[\alpha P_M^* + (1 - \alpha)P_X^*]}{P_N}. \quad (1)$$

The long-run equilibrium implied by this model requires both internal and external balance. Internal balance requires that the nontradable goods market clears, so that

$$C_N(e_M, a) + G_N = Q_N(e_X). \quad (2)$$

Edwards (1988b) uses (2) to show that the equilibrium price of nontradables can be represented as

$$P_N = v(a, g_N, P_M^*, \tau), \text{ where } \frac{\partial v}{\partial a} > 0, \frac{\partial v}{\partial g_N} > 0, \frac{\partial v}{\partial P_M^*} > 0, \frac{\partial v}{\partial \tau} > 0. \quad (3)$$

Internal balance does not require that the current account be zero, but external balance requires that the intertemporal external budget constraint, i.e. the discounted sum of present and future current accounts has to be zero. This is compatible with sustainable capital flows in the long run (i.e. $CA = \dot{R} = 0$) and sustainable fiscal policy (i.e. $\dot{D} = 0$). Edwards uses these considerations and (3) to show that the long-run equilibrium real exchange rate is

$$q_{LR} = v(m_0 + \rho_0 F_0, g_{N_0}, P_{M_0}^*, \tau_0), \quad (4)$$

where a subscript $_0$ denotes a variable at the steady-state level and $m = M/E$. Equation (4) shows that the long-term equilibrium real exchange rate is just a function of real variables, called the *fundamentals*. These fundamentals can shift in both the short and long run, and they are normally influenced by other macroeconomic variables. Further, both real and monetary variables can influence the short-term real exchange rate, since monetary variables (such as D and E) can influence the short-run value of a and hence P_N . The empirical model that we study below defines the real exchange rate similarly to (1) and it incorporates both real and nominal variables.

Edwards studies the short-run and long-run implications of this model via a phase diagram that incorporates an upward sloping $\dot{\rho} = 0$ schedule (driven by $\dot{\rho} = \rho L(m/\rho F)$ with $L'(\cdot) < 0$) and a downwards sloping $\dot{m} = 0$ schedule (driven by $\dot{m} = Q_X(e_X) - C_M(e_M, a) + g_N - t/E$). The fact that ρ reflects the spread between a freely floating and a fixed exchange allows the model to capture changes in exchange rate regimes via shifts in the $\dot{\rho} = 0$ schedule. Although this version of the model does not directly incorporate capital flows, he points out that one could treat these as

an exogenous source of funds for Government and then adjust the intertemporal external budget constraint accordingly. He studies this in more detail in Edwards (1989) and shows that capital inflows and other relaxations of exchange controls lead to increased demand for non-tradable goods and a real exchange rate appreciation.

3.1.1 Real Exchange Rate Response to Trade Liberalization

We can use the above model to examine the effects of trade liberalization on real exchange rates. A trade liberalization generated by a(n) (unanticipated) reduction in import tariffs increases the demand for traded goods via a decrease in e_M , but substitution of tradables for nontradables also causes a decline in the price of nontraded goods, that counters the decline in e_M and increases e_X . The net effect of these changes can shift the $\dot{m} = 0$ schedule in either direction, but assuming that the direct effect of e_M on traded goods is dominant, this will shift the $\dot{m} = 0$ schedule leftwards, causing long-run declines in m and ρ , corresponding declines in real assets a and the demand for nontradables ($C_M(e_M, a)$), and a corresponding long-run depreciation.³

3.1.2 Real Exchange Rate Response to Expansionary Government Consumption

We can also use the above model to analyze the effects of changes in government consumption on the real exchange rate. If we firstly interpret this increase as an increase in government's demand for nontradables g_N , then this will create higher demand and thus a rise in the price of nontradables, generating a real appreciation. However, regardless of its tradable/nontradable composition, government consumption is financed by public debt that must be paid back. This increase in taxes t will lead to a fall in household assets, a corresponding decline in demand for nontradables and a corresponding real depreciation. Given these two channels, the effect of an increase in government consumption is a priori indefinite as it depends on the sum of these two effects - a point that is easily reinforced by noting that the $\dot{m} = 0$ schedule will move outwards with increased g_N , but inwards with increased t . Following Edwards (1989) and assuming that government expenditure is mostly on nontradable goods, we expect the first channel to play the more dominant role, leading to a real appreciation in response to an increase in government consumption.

3.1.3 Real Exchange Rate Response to Traded-Sector Productivity Growth

The well-known explanation of the relationship between the long-run real exchange rate and productivity differentials between traded and nontraded goods is the Balassa (1964) - Samuelson (1964)

³These changes can also be understood by considering the relationship $\frac{dq_{LR}}{d\tau} = -(\frac{q}{P_N})[(\frac{\partial v}{\partial a})(\frac{\partial m}{\partial \tau} + \frac{\partial \rho}{\partial \tau}) + \frac{\partial v}{\partial \tau}]$.

hypothesis. The intuition is that higher productivity growth in the traded sectors (given that productivity growth in the nontraded sector for most countries is slow) induces a rise in wages in the traded sectors, and this raises wages in the nontraded sectors due to labor mobility. The nontraded 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 nontraded goods at home than abroad will induce a long-run appreciation of the domestic currency.

The effects of productivity gains can be analyzed using the Edwards model explained above. For instance, productivity improvements in the traded sector can be captured by a shift in $Q_X(e_X)$ so that $Q_X^{New}(e_X) > Q_X(e_X) \forall e_X$. This shifts the $\dot{m} = 0$ schedule to the right, and ultimately leads to a long-run real appreciation via an increase in wealth. However, the effect of productivity improvements in the nontraded sector is less clear, because the relative supply and demand elasticities associated with nontradables determine whether the price of nontraded goods P_N will increase or decrease, which in turn affects the direction of shifts in the demand for imports $C_M = C_M(e_M, a)$ and the supply of exports $Q_X(e_X)$. Therefore technological progress can cause either a real depreciation or a real appreciation depending on which sectors of the economy are affected and whether supply or demand effects dominate. Edwards (1989) discusses these issues in considerable detail, and concludes that long-run appreciation is more likely than long-run depreciation.

3.1.4 Real Exchange Rate Response to Contractionary Monetary Policy

Dornbusch's (1976) overshooting hypothesis is a baseline theory used to study the effects of monetary policy on exchange rates. The Dornbusch's model is based on the three main assumptions: price stickiness, uncovered interest parity (UIP), and long-run purchasing power parity (PPP). The economy is small, capital is perfectly mobile, and investors have rational expectations. According to this model, an unanticipated increase in domestic interest rates causes domestic assets to become more attractive to investors, inducing net inflows on the capital account and boosting the supply of foreign currencies. As a result, the price of foreign currency falls substantially in the short run and this leads to an initial appreciation of the domestic currency that overshoots its new long-run equilibrium level. However, a subsequent depreciation of the domestic currency in the future is then expected, in line with UIP, and the price level gradually adjusts to the new equilibrium in the long run. Therefore a short-run appreciation beyond its long-run value is followed by a depreciation towards the terminal value to assure UIP and long-run PPP.

4 The Model and Identification Procedure

Our model relates to fourteen developing countries in Eastern and Southern Asia - China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan, Sri Lanka, Vietnam, Brunei, Myanmar, Bangladesh, Nepal and Taiwan - and it contains domestic and foreign variables. Instead of using US variables as a proxy for foreign variables, we consider thirty-two countries/regions when constructing country-specific foreign variables.⁴ Let $g_{i,t}$ be a $k \times 1$ vector of domestic variables, and $g_{i,t}^*$ be a $k^* \times 1$ vector of country-specific foreign variables and global factors (such as oil prices). Each country-specific foreign variable is constructed as a country-specific weighted average of corresponding domestic variables of all other countries/regions giving rise to

$$g_{i,j,t}^* = \sum_{l=1}^{32} w_{i,j,l} g_{l,j,t}, \quad i = 1, 2, \dots, N,$$

where N ($= 14$) is the number of countries in the model and $g_{i,j,t}(g_{i,j,t}^*)$ is the element of $g_{i,t}(g_{i,t}^*)$ corresponding to variable j . Thus, i (and l) index countries and j indexes variables. 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 l to country i . In particular, for each variable j , $w_{i,j,l}$ is the trade share of country l in the total trade (imports + exports) of country i with all of its trade partners, such that $w_{i,j,i} = 0$ and $\sum_{l=1}^{32} w_{i,j,l} = 1$.⁵

4.1 Structural Vector Autoregressive Model

We construct a panel model for the group of Asian developing countries because the use of panel data methods offers improved power with respect to our initial cointegration tests, and also improves the efficiency of both short run and long run 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, which is reasonable given that our country group consists of developing countries. Later, once we have estimated our model, we perform formal tests of weak exogeneity for foreign and global variables and our test results cannot reject this assumption.

⁴In addition to fourteen 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 States, Korea, Japan, Singapore, Brazil, Mexico, Chile, Argentina, South Africa and Turkey. In line with the developing/non-developing country classification provided in the IMF's World Economic Outlook published in April 2012, we do not classify Japan, South Korea or Singapore as developing countries.

⁵Data sources are described in Appendix A, and all country specific data series together with the 32×14 matrix of the trade shares used for constructing the country-specific foreign variables are available via the online data supplement associated with this paper.

We use residual serial correlation test results to choose an augmented vector autoregressive (VARX*) model, and find that four lags of domestic variables and a single lag for foreign variables provide an appropriate dynamic specification for the panel. The resulting VARX*(4,1) model is

$$g_{i,t} = \Phi_1 g_{i,t-1} + \Phi_2 g_{i,t-2} + \Phi_3 g_{i,t-3} + \Phi_4 g_{i,t-4} + \Psi_0 g_{i,t}^* + \Psi_1 g_{i,t-1}^* + u_{i,t}, \quad (5)$$

for $t = 1, 2, \dots, T$ and $i = 1, 2, \dots, N$. The notation is such that Φ_1, Φ_2, Φ_3 and Φ_4 are $k \times k$ matrices of coefficients associated with lagged endogenous/domestic variables, Ψ_0 and Ψ_1 are $k \times k^*$ matrices of coefficients associated with weakly exogenous/foreign variables, and $u_{i,t}$ is a $k \times 1$ 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 and we use an additional dummy to allow for a possible structural shift at the time that the Asian Financial Crisis started (1997). We also include a subset of cross-sectional means to account for cross-country correlation, but we drop these from (5) to simplify our exposition.

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

$$\Delta g_{i,t} = -\Pi z_{i,t-1} - (\Phi_2 + \Phi_3 + \Phi_4)\Delta g_{i,t-1} - (\Phi_3 + \Phi_4)\Delta g_{i,t-2} - \Phi_4\Delta g_{i,t-3} + \Psi_0\Delta g_{i,t}^* + u_{i,t},$$

where $\Pi = (I - \Phi_1 - \Phi_2 - \Phi_3 - \Phi_4, -\Psi_0 - \Psi_1)$ and $z_{i,t-1} = \left(g'_{i,t-1}, g'^*_{i,t-1}\right)'$, (6)

and the cointegrating relationships between the variables are summarized in the $k \times (k + k^*)$ matrix Π . If the rank of Π is $r \leq k$ then there are r long-run relationships between the variables, and the matrix Π can be written as $\Pi = \alpha\beta'$, where α is a $k \times r$ matrix of adjustment coefficients and β is a $(k + k^*) \times r$ matrix that specifies the r long-run relationships. We impose estimated cointegrating relationships $\hat{\beta}$ that are supported by cointegration tests, and use procedures described in Sections 5.2 and 5.3 to estimate

$$\Delta g_{i,t} = B_0 ecm_{i,t-1} + B_1 \Delta g_{i,t-1} + B_2 \Delta g_{i,t-2} + B_3 \Delta g_{i,t-3} + B_4 \Delta g_{i,t}^* + u_{i,t}, \quad \text{where} \quad (7)$$

$$ecm_{i,t-1} = \hat{\beta}' z_{i,t-1}, \quad B_0 = -\alpha, \quad B_1 = -(\Phi_2 + \Phi_3 + \Phi_4), \quad B_2 = -(\Phi_3 + \Phi_4), \quad B_3 = -\Phi_4 \quad \text{and} \quad B_4 = \Psi_0.$$

Our interest is to examine impulse responses to economically meaningful structural shocks associated with the VARX* in (5). The cross-correlation between the k elements of $u_{i,t}$ in (5) prevents the interpretation of any given element in $u_{i,t}$ as a separate structural shock, but we can separately identify structural shocks via a matrix A such that $A^{-1}u_{i,t} = v_{i,t}$ and $E(v_{i,t}v_{i,t}') = I_k$, because the k elements in $v_{i,t}$ are independent. There are many A that can render this independence within $v_{i,t}$ and we choose A in a way that ensures that the elements in $v_{i,t}$ have meaningful economic

interpretations. Treating the k elements in $v_{i,t}$ as fundamental shocks and noting that $u_{i,t} = Av_{i,t}$, we can see that the j th column of A (we call this $a^{(j)}$) measures the impact on the vector $u_{i,t}$ (and hence on $g_{i,t}$) of a unit shock to the j th element of $v_{i,t}$ (assuming no shocks to any of the other elements in $v_{i,t}$).

4.2 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 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 -dimensional vector m of unit length such that $a = \vec{A}m$, where $\vec{A}\vec{A}' = \Sigma$ and \vec{A} is any arbitrary decomposition of Σ such as a lower triangular Cholesky factor,⁶ 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 . Thus

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

where m_j is the j^{th} element of m ; $r_j^c(h) \in \mathbb{R}^k$ is a $k \times 1$ vector of the impulse response at horizon h to the j^{th} shock in a Cholesky decomposition of Σ . 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 .

Moreover, it is interesting to perform variance decompositions or in other words, compute how

⁶The 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.

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 calculated using

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

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* in (7), 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, \vec{A} is computed using a draw of Σ from the posterior. Consequently, we can calculate the candidate impulse vector as $a = \vec{A}m$.

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 standardized 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

$$a = \arg \min_{a=\vec{A}m} \Psi(a) \quad \text{for}$$

$$\Psi(a) = \sum_{s \in l_{r+}} \sum_{h=0}^{H_{re}} ff\left(-\frac{r_{a,s}(h)}{\sigma_s}\right) + \sum_{s \in l_{r-}} \sum_{h=0}^{H_{re}} ff\left(\frac{r_{a,s}(h)}{\sigma_s}\right), \quad (8)$$

where the penalty function suggested by Uhlig (2005) is

$$ff(w) = \begin{cases} ww & \text{if } ww < 0, \\ 100 \times ww & \text{if } ww \geq 0. \end{cases} \quad (9)$$

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 (8), we can identify the first shock $a^{(1)} = \vec{A}m^{(1)}$.

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 $R = [r_{1,s}^c(0), \dots, r_{k,s}^c(0)]$. Therefore, additionally imposing orthogonality conditions and a zero restriction, we minimize the problem below:

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

For computation, we find a "best impulse matrix" by undertaking the numerical minimization of the above criterion function $\Psi(a)$ in equation (8) on the unit sphere, given each draw of B and Σ . We parameterize the space of unit-length vectors by using stereo projection,⁷ 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 different by less than 0.01, we keep the impulse vector. In contrast, if the two minima differ by more than 0.01, we keep only the vector which generates the smaller value of the total penalty, and 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 16th, 50th and 84th percent quantiles for each h -step-ahead forecast.

⁷The 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.

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 fourteen developing countries in Asia - China, Malaysia, Indonesia, Philippines, Thailand, India, Pakistan, Sri Lanka, Vietnam, Brunei, Myanmar, Bangladesh, Nepal and Taiwan. 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 includes the log traded-nontraded productivity differential ($x_{i,t}$);⁸ 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.⁹

When constructing the sectoral productivity differential ($x_{i,t}$), it is important to classify economic activities into traded and nontraded sectors. This is a weak aspect many previous studies of the BS hypothesis. We address this by using the classification introduced by Dumrongrittikul (2012). This classification 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 traded in international markets, so that it is appropriate to assess the tradability of each industry using international trade data from input-output tables. The second concept is that the price of traded goods and services is more likely to follow PPP and the law of one price than the price of nontraded goods and services. This is examined by using econometric methods to test the comovement of prices in each industry with world prices, and the results are used to aid classification.

Edwards (1988b, 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)

⁸Labour productivity is used as a proxy for total productivity due to data limitations. Our measure of productivity is the ratio of value added in constant local currency units to the number of employees in each sector.

⁹It is difficult to find a good proxy for trade liberalization due to unavailability of long-period data. However, this ratio has been used elsewhere in the literature as a measure for trade liberalization and tariff reduction (examples include Elbadawi (1994), and Candelon et al (2007)). While this measure can be affected by other more general phenomena such as globalization, it is primarily driven by trade policies.

and the inflation rate ($\pi_{i,t} = p_{i,t} - p_{i,t-1}$). The model includes foreign analogues of short and long run exchange rate determinants, and the log oil price index (oil_t) that we use to proxy for unobserved global factors.¹⁰ Data sources are provided in Appendix A, as well as plots of the real exchange rates and their long run determinants in Figures A1 to A5.

The real exchange rate plots show considerable variation in behavior across countries and time, reflecting changing country specific (and often nominal) exchange rate policies, as well as different price pressures across countries and time. On average, real exchange rates have depreciated in most countries over the years in the sample, although Vietnam and Myanmar provide exceptions to this pattern and most countries have experienced several changes between episodes of depreciation and appreciation. Real GDP has risen for all countries, as has openness (except for Myanmar and Brunei). Traded non-traded productivity has risen in many countries, but since this variable depends on country specific composition of output which changes over time, we see considerable variation in this measure, both across countries and time. It is clear that real output and productivity differentials are not strongly correlated, so that each has the potential to contribute to real exchange rates in different ways. Patterns in real government spending are country specific, and most do not tend to trend upwards or downwards for long periods of time.

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 ($\pi_{i,t}$).

5.2 Testing for the Long-Run Relationships

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

$$\text{Purchasing Power Parity} \qquad q_{i,t} \sim I(0) \quad (\text{R1})$$

$$\text{Fisher Equation} \qquad si_{i,t} - \pi_{i,t} \sim I(0) \quad (\text{R2})$$

$$\text{Output Convergence} \qquad y_{i,t} - y_{i,t}^* \sim I(0) \quad (\text{R3})$$

$$\text{Uncovered Interest Parity} \qquad si_{i,t} - si_{i,t}^* - E(\Delta e_{i,t+1}) \sim I(0) \quad (\text{R4})$$

¹⁰Some authors (eg Chinn (2000)) have treated the oil price index as a long-run determinant of the real exchange. In our setting, we found that although cointegration restrictions that involved the oil price were weakly supported by cointegration tests, the imposition of these restrictions led to models that exhibited strong evidence of dynamic misspecification.

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

$$q_{i,t} - \lambda_1(x_{i,t}^* - x_{i,t}) - \lambda_2(y_{i,t}^* - y_{i,t}) - \lambda_3(gov_{i,t}^* - gov_{i,t}) - \lambda_4(tt_{i,t}) - \lambda_5(open_{i,t}) \sim I(0) \quad (R5)$$

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.¹¹ 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 suggested by Im et. al. (2003) and Maddala and Wu (1999) on $q_{i,t}$, $si_{i,t}$, $y_{i,t} - y_{i,t}^*$ and $si_{i,t} - si_{i,t}^*$ to check the validity of the relationships (R1)-(R4). The results show that $q_{i,t}$, $si_{i,t}$ and $y_{i,t} - y_{i,t}^*$ are approximately $I(1)$, suggesting that PPP, the Fisher Equation and output convergence do not hold. However, the test statistics seem to suggest that $si_{i,t} - si_{i,t}^*$ is stationary. This provides strong evidence for UIP in line with Bjørnland (2009) and what we expect, as nowadays financial markets in most countries are more closely linked with each other, and have become more like a global financial market.

We apply four residual-based cointegration tests suggested by Pedroni (1999) to test for a long-run relationship that conforms with (R5).¹² 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. These tests allow cointegrating vectors to be different for each country, and since they are based on the assumption of cross-

¹¹Actually, the Solow-Swan model implies convergence of output per capita, not output. Convergence of output per capita will occur in the Solow-Swan model only if the rate of technological progress converges across countries.

¹²This testing procedure differs from the above because it involves estimating the potentially cointegrating vector. Wagner and Hlouskova (2010) show that Pedroni's (1999) tests perform well compared to other panel cointegration tests. We focus on four out of the seven test statistics suggested in Pedroni (1999) because his subsequent work in 2004 shows that the panel and group t-statistics have higher power than the other three test statistics in a situation similar to ours.

sectional independence in errors, we include a set of common time dummies in the hypothesized cointegrating regression to accommodate cross-sectional dependence across countries.

Table 1: Cointegration tests for the relationship (R5)

Variables	Panel PP	Group PP	Panel ADF	Group ADF
$q_{i,t}, (x_{i,t}^* - x_{i,t}), (y_{i,t}^* - y_{i,t}),$ $(gov_{i,t}^* - gov_{i,t}), tt_{i,t}, open_{i,t}$	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 numbers in the tables report the P-values for tests of the null hypothesis of no cointegration.

2) The lag length in ADF-type regressions used in Pedroni's tests is selected by AIC with a maximum of 4 lags.

Table 1 reports the results of the panel and group t-statistics of Pedroni's cointegrating tests. The evidence does not support relationship (R5). However, as (R5) relates to a large set of variables and the span of our data set for each country is quite 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 four fundamentals i.e. $(x_{i,t}^* - x_{i,t}), (y_{i,t}^* - y_{i,t}), (gov_{i,t}^* - gov_{i,t}),$ and $open_{i,t}$, but do not find evidence that the terms of trade contributes to this relationship.

5.3 Estimating and Interpreting the Model

5.3.1 Estimating and interpreting the cointegrating vector

We use within-dimension panel DOLS estimation developed by Mark and Sul (2003) to estimate (R5), which measures the long-run (cointegrating) relationship between the real exchange rate and its fundamentals.¹³ The within-dimension estimator is somewhat restrictive since it assumes that the cointegrating vector is the same across cross-sectional units. However, our countries are broadly similar in the sense that they are all developing countries within the same region that experience similar influences from the developed world, and our use of the panel allows more precise point estimates of the cointegrating vectors. We allow for heterogeneity across countries through our use of country-specific short-run dynamics and fixed effects. The estimated DOLS equation for the real exchange rate relationship is

$$\hat{q}_{i,t} = d_i + \theta_t - 0.026(x_{i,t}^* - x_{i,t}) - 0.375(y_{i,t}^* - y_{i,t}) + 0.149(gov_{i,t}^* - gov_{i,t}) + 0.303open_{i,t} + \delta_i' sd_{i,t},$$

(0.048) (0.065) (0.076) (0.041) (10)

where $i = 1, 2, \dots, N$; $sd_{i,t} = (\Delta f d'_{i,t-1}, \Delta f d'_{i,t}, \Delta f d'_{i,t+1})'$ where $f d_{i,t} = ((x_{i,t}^* - x_{i,t}), (y_{i,t}^* - y_{i,t}), (gov_{i,t}^* - gov_{i,t}), open_{i,t})'$; d_i is a country-specific effect; θ_t is a common time-specific factor which is used

¹³This 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 (see Kao and Chiang (2000)).

to capture some forms of cross-sectional dependence across countries¹⁴. 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.

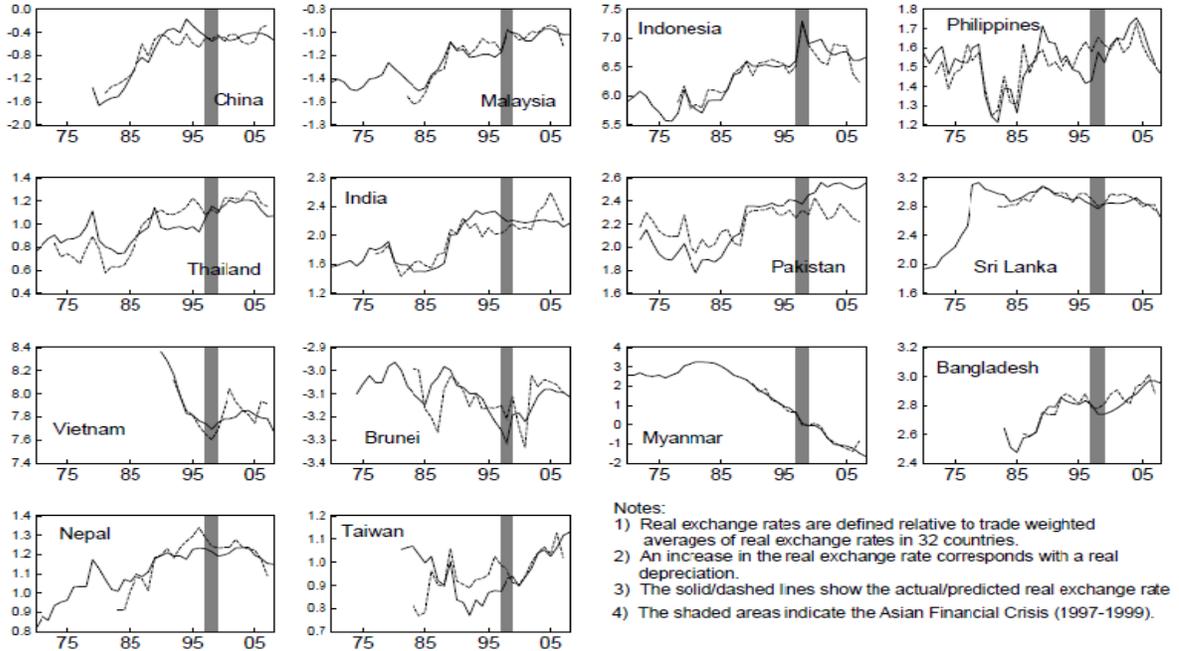
The signs of those coefficients that directly reflect long-run policy influences on real exchange rates correspond to our expectation i.e. after controlling for other factors, an increase in the degree of openness in the economy is associated with a real depreciation and expansionary government consumption is associated with a real appreciation. However, the coefficient estimates associated with traded sector productivity gains or total gains in output (relative to foreign economies) have unexpected signs, suggesting that such gains are associated with a real depreciation, after controlling for other factors. A possible explanation for the first of these unexpected effects (which is small and not even statistically significant) is that increases in $(x_{i,t}^* - x_{i,t})$ might be primarily generated by productivity growth in the non-traded sector of the domestic economy, which has unclear effects on the prices of non-traded goods and hence on the real level of exchange, as discussed in Section 3.1.3 above. Further, if increases in the output differential $(y_{i,t}^* - y_{i,t})$ are mostly due to growth in foreign output, (rather than a decline in domestic output) and this triggers demand for exports from Asian to foreign countries, then a negative sign (corresponding to an appreciation of Asian exchange rates) would be expected.

Figure 1 plots the observed real exchange rates together with those predicted by country fundamentals (using (10)). The exchange rate is overvalued relative to fundamentals when the solid line is under the dashed line. While it is clear that the real exchange rates for most countries follow paths that are largely in line with fundamentals, each country has also experienced periods when its real exchange rate has been misaligned. The shaded areas on the graph relate to the Asian Financial Crisis which ran from 1997 to 1999, and we can see that the exchange rates for Malaysia, Philippines, Indonesia and Thailand were all overvalued in the years leading up to this crisis. Another interesting observation is that while fundamentals suggest that the Chinese exchange was undervalued in the late 1990's, it has not been systematically undervalued since that time, contrary to claims made by policy makers in foreign developed economies¹⁵.

¹⁴Note that θ_t implicitly allows for common structural shifts and such a shift was observed over 1981-1983 (but not over the 1997-1999 Asian financial crisis period). We include an Asian financial crisis shift dummy in the VECMX that is studied below to account for the possible effects of this crisis.

¹⁵See, for instance, Cooper (2006) and Morrison and Labonte (2013).

Figure 1: Real exchange rates and those predicted in long-run equilibrium



5.3.2 Estimating and interpreting the VECMX

The VECMX* in (6) contains seven endogenous variables and six weakly exogenous variables, i.e.¹⁶

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

$$g_{i,t}^* = [x_{i,t}^*, y_{i,t}^*, gov_{i,t}^*, si_{i,t}^*, \pi_{i,t}^*, oilt]'$$

and our estimation allows for fixed effects and a dummy to account for possible change after the Asian Financial Crisis. Note that when we examine the long-run relationship (R5) in Section 5.2.1 most variables are in differentials (differences between domestic and foreign variables), but when we estimate (6), we impose the estimated cointegrating relationship in R5, but separate the elements of z_{it} into domestic and foreign variables. The VECMX incorporates two long run relationships (i.e. the exchange rate relationship discussed above and uncovered interest rate parity) together with the inflation rate as error correction terms.¹⁷ We follow Johansen (1992) and conduct formal tests that our foreign and global variables are weakly exogenous, and find no evidence to reject weak exogeneity. See Table B1 in Appendix B.

While we base most of our interpretation on the impulse response that follows, we comment on

¹⁶Note 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.

¹⁷We treat inflation analogously to the error correction terms because it is stationary. The estimated model also includes cross country means of endogenous variables to control for unobserved global common factors.

some of the VECM coefficient estimates first. The estimated adjustment coefficients are tabulated in Table B2, and we note that the estimated adjustment coefficient in the exchange rate equation that relates to (lagged) deviations of the real exchange rate from long-run equilibrium is -0.0587 . On its own, this would imply that the half-life associated with exchange rate misalignment is 11.5 years and long (even for developing economies), but our model also implies that government spending adjusts when the exchange is misaligned, and this also contributes to realignment. In contrast, the estimated adjustment coefficient in the interest rate equation that relates to (lagged) deviations from uncovered interest rate parity is -0.4026 , which implies a shorter half-life of 1.34 years. High domestic inflation leads to appreciation and less international trade, but it is not sustainable, with a half life of only about 8 months. We do not tabulate other coefficients associated with VECM given space considerations, but note in passing that oil price increases do not have a statistically significant effect on exchange rate movements. However, they have a positive and strong impact on growth, consistent with the effects of global aggregate demand shocks, as discussed in Killian (2009) and Baumeister et al (2012). The relevant coefficient errors and standard errors are $-.0264$ ($s.e. = 0.0248$) and 0.0861 ($s.e. = 0.0141$). The shift dummy that we used to account for possible change after the Asian crisis is not statistically significant in the exchange rate equation, but it indicates that openness fell in the wake of the crisis while domestic interest rates rose.

5.4 Shock Identification Assumptions

Table 2 summarizes the set of restrictions adopted in this paper.

Table 2: 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 domestic prices. 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. This type of productivity shock is central to the BS hypothesis, and an associated increase in output is expected. In our case, the restriction that output must rise serves an additional purpose, because it ensures that the positive shocks to the

traded-non-traded productivity differential that we consider arise from positive shocks to traded sector productivity rather than from negative productivity shocks to the non-traded sector. The government consumption shock is characterized 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 nontraded sectors is assumed to be initially unchanged. These restrictions are consistent with the standard New Keynesian model and empirical work in Peersman (2005) and Farrant and Peersman (2006). The zero restriction on the impact response of the productivity differential between traded and nontraded sectors is plausible because an unexpected change in monetary policy should influence productivity in both sectors in the same way on impact.

In short, these restrictions are in line with the theoretical and empirical literature and are sufficient to uniquely disentangle the shocks of interest. 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 interaction of real economic activity between countries. Then we choose the shock to traded-sector productivity as our second shock for identification because productivity will take time to adjust, similar to the international trade share. We check that our results are robust to change in the order of these first two shocks in Section 5.8. 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.

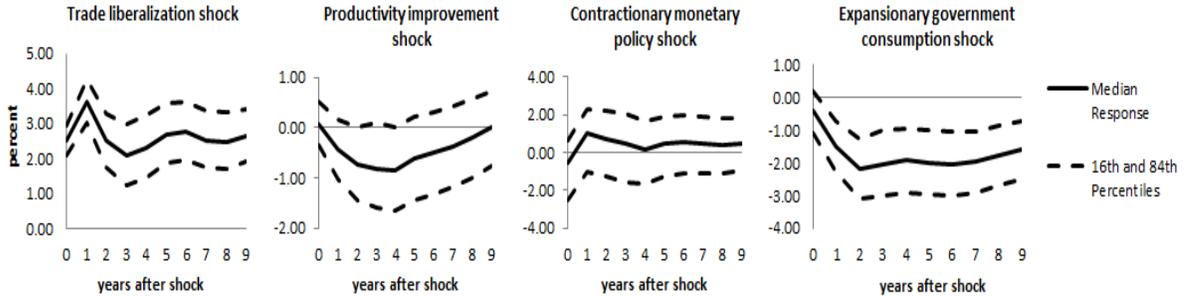
5.5 Impulse Response Analysis

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.

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 2 shows the median impulse responses of the real exchange rate together

with the 16th and 84th percentile responses over nine years after the shocks. Recall that the exchange rate is defined so that it declines as the real value of the home country currency increases. Note that the impulse responses of other variables are provided in Figures C1-C4, Appendix C.

Figure 2: Impulse responses of real exchange rates to each shock of one standard deviation in size.



Note: The four shocks imply changes of approximately 4.8%, 8.4%, 1.8% and 5.9% in the four policy variables.

5.5.1 Trade Liberalization Shock

As shown in Figure 2, our analysis provides evidence that trade liberalization results in a long-run real depreciation of the domestic currencies, in line with Li (2004). This suggests that trade liberalization results in substitution in demand away from nontradables and into importables, and induces a lower value of real assets. This in turn entails a decline in demand and price of nontradables, implying a real depreciation. This is the type of reaction of the equilibrium real exchange rate that, according to Edwards and a variety of theoretical models,¹⁸ is the most likely to occur. We also briefly note that Figure C1 in the Appendix shows that the trade liberalization shock increases real output, in line with theory that predicts gains from trade.

5.5.2 Traded-Sector Productivity Improvement Shock

Our results show that a productivity shock causes a short-run appreciation as predicted by the BS hypothesis, but we find little evidence that this effect persists over a longer horizon. This is in line with the commonly made observation that real exchange rates in countries such as China and India have not experienced long-run appreciation, even though multilateral firms have been present in these countries for quite some time.

Our results differ from the traditional BS model and the related empirical literature for two main reasons. The first reason is that the traded-nontraded productivity differential is only one of several

¹⁸Dornbusch (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.

factors that can cause a persistent deviation of the real exchange rate from its equilibrium. In our model we control for other factors that may have influenced the real exchange rate, so that the measured effect of productivity differentials on the real exchange rate may be lower after capturing these factors. Put differently, the estimated productivity effect based on a simple two-variable model may pick up the effects of other factors on the behavior of the real exchange rate.

The second reason is that previously used measures of sectoral productivity differentials may have been too blunt. One common shortcoming of these measures is that the classification of industries into traded and nontraded sectors in all countries is the same and does not change across time. Edwards (1989) mentions this issue in his empirical work after he found a result that contradicted the BS hypothesis. Richi et al (2013) find that a productivity differential shock causes a small long-run appreciation when they use a measure based on sectoral labor productivity in a detailed breakdown of industries, but their classifications across traded and nontraded industries are the same for all forty eight countries that they consider. Our use of the Dumrongrattikul (2012) classification incorporates a detailed breakdown of industries and also allows changes in the tradable vs nontradable classifications of industries across countries and periods. While other classifications usually treat agricultural goods as traded for all countries, our classification of these goods as nontraded in most Asian countries reflects the fact that most agricultural produce within Asia is domestically consumed. Further, this classification incorporates several instances of classifications of manufacturing and/or trade sectors for China, India and Thailand, which change from nontraded to traded in the mid 2000s - in line with developmental trends observed in these countries.

We find that a productivity differential shock only causes a short run appreciation in Asia.¹⁹ This might be because some of the BS assumptions are not satisfied in Asia, particularly the labor arbitrage condition. We note that although industrial productivity seems to be higher than rural productivity in most Asian developing countries, rural workers often move to nontraded service sectors when they move, rather than move into industrial sectors (that are typically traded). Therefore, traded-sector productivity gains need not raise wages and prices in nontraded sectors.

5.5.3 Contractionary Monetary Policy Shock

Dornbusch's (1976) overshooting hypothesis predicts an immediate appreciation, followed by depreciation that satisfies UIP and ensures that the policy has no long-run effect on the real exchange, as discussed in Section 3.1.4. We see this pattern in the third portion of Figure 2, but note that this

¹⁹We run a simple regression of the real exchange rate on sectoral productivity differential in levels using fixed effects estimation, and 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 long-run movement of the real exchange rate.

support for the overshooting hypothesis is quite weak given the width of our uncertainty bands. We see no evidence of an exchange rate puzzle (immediate depreciation or delayed appreciation, as discussed by Eichengreen and Evans (1995) or Scholl and Uhlig (2008)), despite the fact that our sign restrictions do not rule this possibility out. The clear decline in the inflation response in Figure C3 shows that our sign restriction on inflation has ruled out the well known price puzzle (a positive inflation response to contractionary monetary policy, first pointed out by Sims in 1992). Our sign restrictions also rule out an output puzzle (a positive output response to contractionary monetary policy) in the short-run, but we note that neither the output nor productivity responses are statistically strong in the long-run, consistent with theoretical long-run neutrality of money.

5.5.4 Expansionary Government Consumption Shock

Edwards (1988b and 1989) shows that a rise in government consumption of nontradables will lead to a long-run real exchange rate appreciation as it increases the demand and price of nontradables. This idea contradicts Ricardian equivalence, which asserts that government consumption will not affect aggregate demand in the economy because higher government demand is offset by a reduction of private demand due to their anticipation of higher taxes at some time in the future. This implies no change in the real exchange rate after the shock.

Our finding shows that the shock to government consumption does matter for real exchange rates, inconsistent with Ricardian equivalence. The responses show 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 and the existing literature, e.g. Chinn (1999) and Galstyan and Lane (2009). This suggests that government most likely allocates its consumption expenditure to nontraded goods and services, and the revenue effect is relatively small.

5.6 Forecast Error Variance Decomposition Analysis

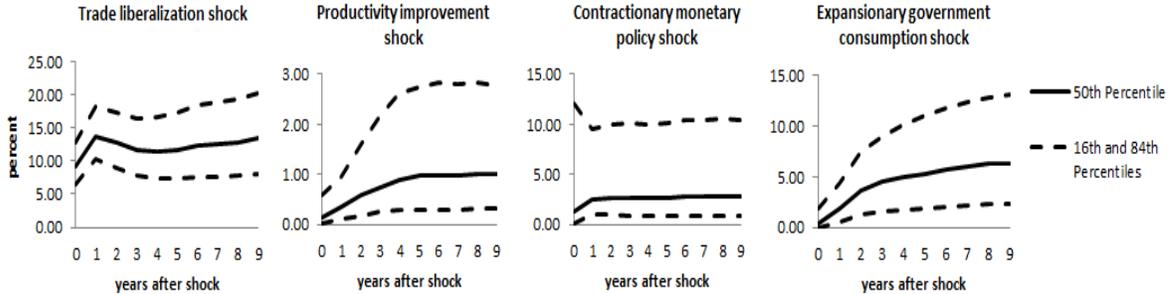
This analysis answers the question of how much of the forecast error variance in real exchange rates can be explained by the four shocks. Table B3 in Appendix B show that these four shocks together can explain a moderate proportion of the total forecast error variance of real exchange rates (i.e. about 24%), in both the short and the long run.

Figure 3 shows how the contributions of each of our four types of shocks to forecast error variance change with the forecast horizon.²⁰ We note that trade liberalization accounts for most of our identified forecast error variance share (i.e. about 83 percent), with its influence rising during the first year, peaking during the second and then dropping down slightly and maintaining a fairly

²⁰Note that the individual graphs have different scales.

constant effect thereafter. This is consistent with our expectation that trade reforms will have a direct and lasting effect on real exchange rates.

Figure 3: Shares of forecast error variance of real exchange rates attributable to each shock.



The initial forecast error variance shares explained by productivity improvement, monetary policy and government consumption shocks are all much smaller, but the relative influence of government consumption shocks grow for a few years as the longer run effects of the shock induce further real appreciation, and eventually this contribution settles at about 6 percent of the total forecast error variance. The effects of productivity shocks follow the same pattern as government consumptions shocks, but they are smaller throughout and eventually contribute a mere 1% percent towards the total forecast error variance. The share due to monetary policy shocks remains virtually unchanged after the one year horizon, contributing only about 3% towards total forecast variance in the short run, with no further contributions thereafter.

5.7 Robustness Analysis

We perform five types of robustness checks to assess whether our empirical results are sensitive to various aspects of our modelling procedures. First, we re-estimate the model using three different subsets of countries - the country group without China, and the country group without India, and the country group without Indonesia - to check whether heterogeneity among these fourteen Asian developing countries might affect our results. Second, we assess the sensitivity of our model to the Asian financial crisis by re-estimating it over two sub-samples that relate to 1970-1996 and 1997-2008. Third, we examine the sensitivity of our results to the use of an augmented $VECMX^*(4, 0)$ rather than an augmented $VECMX^*(3, 0)$ specification. Fourth, we impose alternative and equally plausible identifying restrictions by (i) dropping the restriction in Table 2 that inflation will rise after expansionary government consumption shocks (to obtain identification scheme 1); and (ii) imposing an additional restriction (relative to those in Table 2) that trade liberalization shocks will cause real output to increase (this leads to identification scheme 2). Finally, we estimate the

model in levels, noting that this implicitly accounts for unrestricted versions of the cointegrating relationships (see Sims et al, 1990).

Figure C5 in Appendix C provides real exchange rate responses to the four shocks, given the above described variations to model specification. Impulse responses are generally similar to the benchmark impulse responses with respect to sign and shape, and their error bands mostly overlap with the benchmark error bands, suggesting that apparent differences with respect to the length of responses are not statistically significant. Thus, our results seem to be robust with respect to heterogeneity among countries, the timing of shocks, choice of lag length, alternative identification schemes, and our imposition of cointegrating restrictions. Nevertheless, two observations deserve additional comment. The first is that although the pre and post crisis responses fall within very similar uncertainty bands as the benchmark bands, the pre-crisis responses seem somewhat smoother, whereas the post crisis responses seem to show more variation. This might signal a potential shift in responses after the Asian financial crisis, although the length of data in each sub-sample is too short to provide strong evidence of such a shift. The second difference in the responses is that although the signs and shapes associated with the responses from modelling in levels are very similar to those associated with the error correction benchmark, the response bands for government consumption shocks in Figure 5C (9) cover zero for most periods. Since the model in levels relaxes cointegration restrictions and therefore decreases the efficiency of parameter estimation, we interpret the change in placement of the error bands as a mere symptom of lower estimation efficiency, when cointegration is not imposed.

5.8 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 and no penalty functions are used. First, we use the set of restrictions in Table 2 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. The additional shocks that we identify are demand, supply and pure exchange rate shocks, and we use the set of sign restrictions suggested by Farrant and Peersman (2006) for this purpose.

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 earlier 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 system penalty function is

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

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 2.

Table B4 in Appendix B summarizes the impulse responses obtained from these experiments and compares them with responses when different identification strategies are used. The first line of each panel summarizes the benchmark responses discussed in Sections 5.4 and 5.5 above, while the second line summarizes the responses when the trade liberalization and productivity shocks are reordered. Comparing these lines, we see that the order of identification has minor effects on responses to openness and productivity shocks, but no effect on the other responses. Turning now to the pure-sign-restriction and the system-penalty-function approaches (lines 3 to 5 in each panel), we can see that if we discount the responses that have been under the influence of identification restrictions (those marked with a dagger), then most of the remaining responses are insignificant as zero falls within the 16th and 84th percentile responses. These weaker responses contrast with our baseline findings, suggesting that the penalty function (8) that we used when constructing our baseline responses provides an effective way of clearly identifying the shocks of interest. Finally, when we compare lines 4 and 5 in each panel, we see that the responses associated with the pure-sign-restriction approach for identifying four (rather than all seven) shocks are very similar, supporting Uhlig's (2005) comment that we can concentrate on only identifying the shocks of interest.

Overall, these results give us more confidence in our identification. The pure-sign-restriction and the system-penalty-function approaches generate just a few significant results, because they relate to only a small set of the restrictions. Thus, at any point in time, there are other 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 shocks of interest. We impose more restrictions using the penalty functions to dampen 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 by successfully narrowing down the range of admissible responses. This resolves some of the inherent ambiguities associated with other approaches.

6 Conclusion

Recent economic and population growth in China, India and other developing countries in Eastern and Southern Asia has led to world-wide recognition that these countries are important participants in the global economy, and that their exchange rate policies matter, not only for Asia, but for the rest of the world. This paper provides detailed empirical analysis of exchange rate movements for developing Asian countries over the period 1970-2008, focussing on countries in Eastern and Southern Asia because these countries are strengthening their ties with the developed world, and they are also becoming more integrated among themselves. We investigate the short-run and long-run effects of trade liberalization, a productivity improvement in the traded sector, contractionary monetary policy and expansionary government consumption on real exchange rates in these developing countries, placing an emphasis on the size and dynamics of exchange rate responses to different types of policy shocks.

Our analysis is based on sign restricted impulse response associated with a panel vector error correction model, and it offers a new perspective on exchange rate movements in developing Asia, because previous related research has had to rely on much shorter time series for these countries and hence work with a restricted set of modelling techniques. Additional aspects of our work that differentiate it from other studies of exchange rate movements is that we augment our panel vector error correction model with country specific foreign variables (instead of treating the USA as "the foreign country"), and we also incorporate the Balassa-Samuelson effect by constructing a traded-nontraded productivity differential due to Dumrongittikul (2012) that depends on country specific trade patterns that vary over time.

Our finding is that real GDP growth, productivity in the traded sector, government consumption share and the degree of openness in the economy have a long-run relationship with the real exchange rate. Further, 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 particularly important for explaining the short-run and long-run dynamics of real exchange rates. Second, the shock on the traded-nontraded productivity differential causes an impact appreciation, although the effect of the shock on real exchange rates is small and short-lived. This is relevant for deliberations regarding real exchange rate behavior among fast-growing countries in Asia, as policy makers have observed that their currencies do not tend to appreciate over the long run, despite strong productivity improvements in the traded sector. Third, a contractionary monetary policy shock has no significant long-run impact on real exchange rates, consistent with

the long-run neutrality of money. Fourth, a rise in government consumption generates a permanent appreciation as expected.

Previous experience has suggested that substantial real appreciation in Asian emerging markets can trigger a financial crisis. Therefore, it is important that policy makers monitor exchange rate fluctuations and aim to keep real exchange rates on a steady or slow moving course. Based purely on our findings, we make the following observations on real exchange rate policy. First, policymakers need to be cautious when implementing a policy based on government consumption or trade, as changes in these policies will have a permanent effect on real exchange rates. Second, our evidence shows that policies designed to increase the openness ratio, interpreted as trade policies in this setting, can have a strong effect on real exchange rates. This suggests that perceived undervaluation of Asian developing country real exchange rates might actually be a natural consequence of trade liberalization, and that attempts to induce real appreciation would not necessarily be appropriate. Additionally, we observe that although a policy related to traded-sector productivity might cause real appreciation in theory, such a policy is likely to be ineffective for managing real exchange rate behavior in Asian developing countries in practice.

The countries in our analysis are broadly similar in the sense that they are all developing countries from the same geographical region, and they experience similar influences from the outside world. Most importantly, these countries are at similar points on their development trajectories, and they are facing the same set of issues with respect to their exchange rate policies. These similarities, together with the fact that these countries are becoming increasingly interrelated provide rationale for analyzing their exchange rates within a panel. Our results would be useful for understanding exchange rate movements in other countries in Eastern and Southern Asia, and they might also aid discussions relating to possibility of currency co-ordination in Asia (see, eg Gupta, 2012). However, we need to point out that our results are not necessarily applicable to countries in Western or Central Asia, because these latter countries are working with different resource structures, different political environments and their development strategies are vastly different. It is also probable that if we *had* included them in our analysis then this would have lowered the homogeneity (and hence usefulness) of the panel. The same caveats apply to countries such as Brazil or Russia, that are developing rapidly but have different interactions with Europe and North America. Naturally, the study of exchange rate movements for a broader set of countries would be useful for purposes that go beyond the scope of this current paper. We are currently developing a global vector autoregression (see Pesaran et al. 2004) to achieve this broader objective and will report on our findings in due course.

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A Data Appendix

The analysis is based on an unbalanced panel of annual time series relating to fourteen developing Asian countries - China, Malaysia, Indonesia, Philippines, India, Pakistan, Sri Lanka, Vietnam, Brunei, Myanmar, Bangladesh, Nepal and Taiwan. The country specific time series include real exchange rates, real GDP, traded-nontraded productivity differentials, nominal interest rates, openness, real government expenditure and terms of trade. Most of these time series span the period from 1970 to 2008 (i.e. 39 observations), although a few series in the panel (most of which relate to Vietnam) start as late as 1990. The baseline panel VECM is estimated using an effective sample of 347 observations.

The primary data sources are listed below, and transformations used to construct the data in the regressions are described in the text and online appendix. Consumer price indices, nominal exchange rates, export value indices, import value indices, government consumption (% of GDP), as well as exports and imports (% of GDP) for all countries except Taiwan were drawn from the World Development Indicators (WDI), although the CPI for Vietnam and the export and import value indices for Thailand, India, and Pakistan were taken from the IMF's International Financial Statistics (IFS), and the government consumption for Myanmar, was taken from Asian Development Bank (ADB) database. All series for Taiwan were taken from Taiwan's national statistics. The IFS money market rate series were used as the short-term interest rate, except for China, Myanmar and Brunei, for which the IFS deposit rate was used, and for Nepal, for which the treasury bill rate was used. We used GDP series measured in constant 1990 local currency units, classified by economic activity (ISIC 3) into seven categories. These were mainly taken from the National Accounts Main Aggregates database, compiled by The United Nations (UN). Employment classified by ISIC 3 was from the LABORSTA database. Oil prices (averages of Brent Crude series) were drawn from Datastream and gold prices were gold fixing prices in London Bullion Market, based in U.S. dollars per troy ounce and drawn from the Federal Reserve Economic data of St. Louis.

Figures A1 to A5 provide plots of the main series of interest.

Figure A1 Real exchange rates

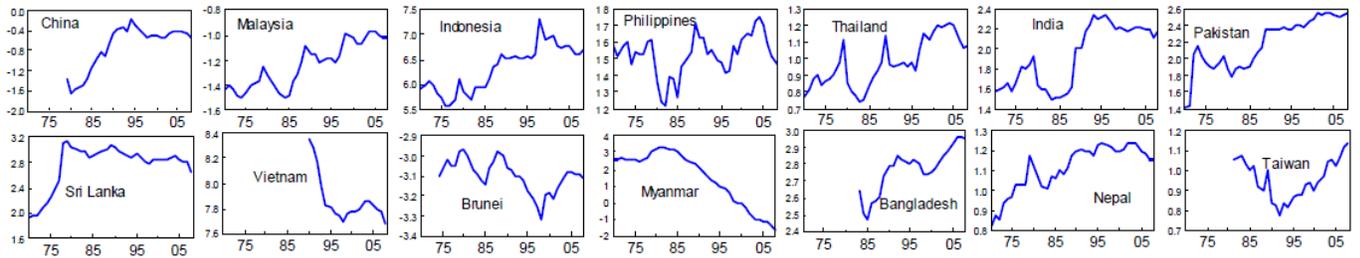


Figure A2 Real GDP

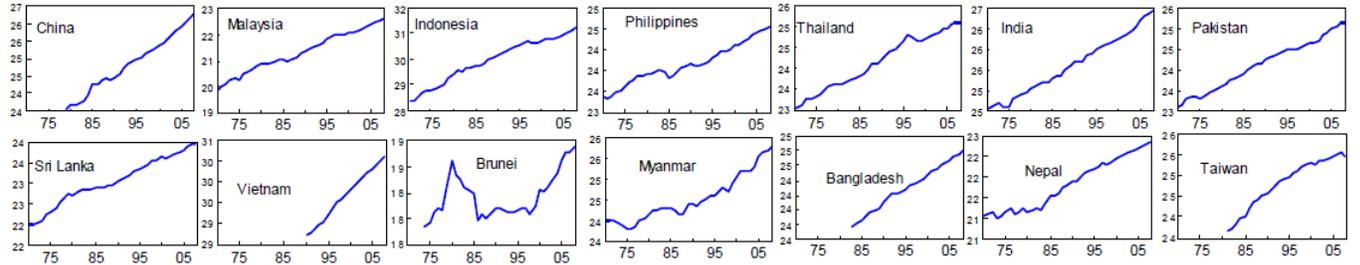


Figure A3 Traded nontraded productivity

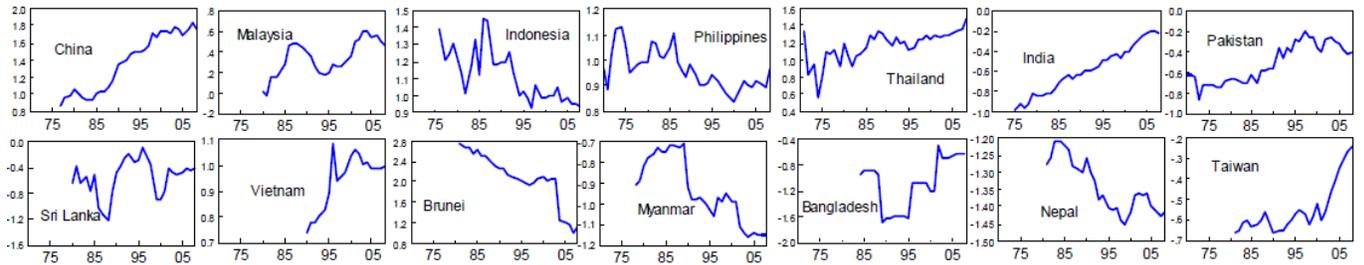


Figure A4: Openness

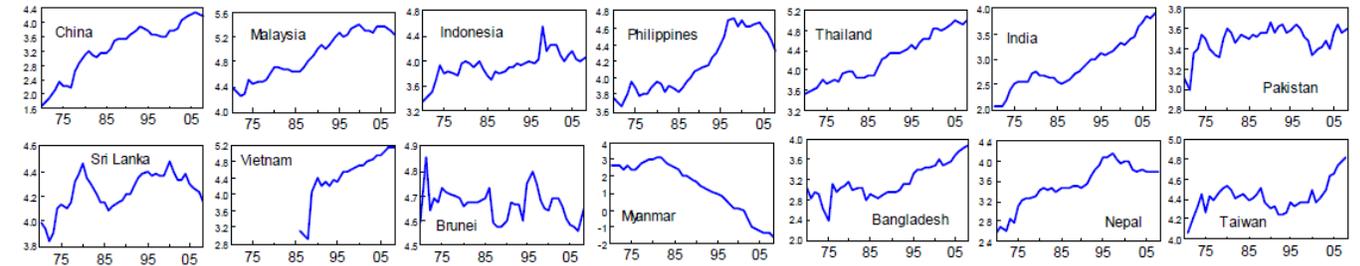
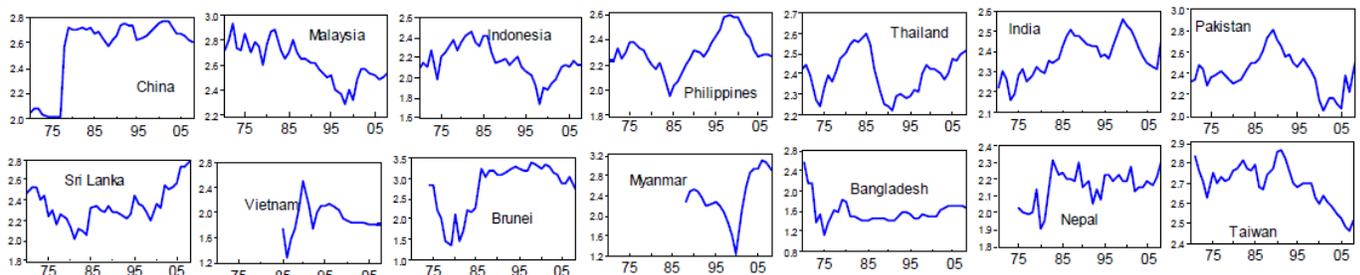


Figure A5: Real government spending



B Table Appendix

Table B1: F statistics for testing the weak exogeneity of the foreign and global (oil price) variables

	$x_{i,t}^*$	$y_{i,t}^*$	$gov_{i,t}^*$	$si_{i,t}^*$	$inf_{i,t}^*$	oil_t
F statistic	0.700	0.828	0.790	1.176	2.139	0.394
(p-value)	(0.553)	(0.479)	(0.500)	(0.319)	(0.095)	(0.757)

Notes: The weak exogeneity test can be performed by estimating

$$\Delta g_{s,t}^* = d_i + \sum_{r=1}^3 \phi_{s,r} ecm_{t-1}^r + \sum_{b=1}^B \theta_{s,b} \Delta g_{t-b} + \sum_{l=1}^L \theta_{s,l} \Delta g_{t-l}^* + \varepsilon_{s,t},$$

where $g_{s,t}^*$ is the s^{th} element of the foreign variable vector g_t^* , d_i for $i = 1, 2, \dots, N$ are country-specific dummies, $\varepsilon_{s,t}$ is an error term and ecm_{t-1}^r are the estimated error correction terms corresponding to the r cointegrating relationships. We choose the lag orders, B and L in the light of serial correlation tests, allowing for a maximum of four lags for both domestic and foreign variables. We test the null hypothesis that the foreign variable is weakly exogenous by testing the joint significance of the estimated error correction terms in the above regression.

Table B2: Estimated adjustment coefficients in the VECM

ECT\Eqn	$\Delta(q_t)$	$\Delta(open_t)$	$\Delta(x_t)$	$\Delta(y_t)$	$\Delta(gov_t)$	$\Delta(inf_t)$	$\Delta(si_t)$
$R(5)_{t-1}^\dagger$	-0.0587	0.0242	.0060	.0102	-0.0754	.0186	-.004
(st-error)	0.021	0.019	0.029	0.012	0.027	0.010	0.007
$si_{t-1} - si_{t-1}$.0470	.2659	-.2980	-.0182	0.0513	-.1014	-.4026
(s-error)	0.168	0.152	0.228	0.095	0.213	0.082	0.059
inf_{t-1}	-0.5564	-0.6397	.0745	-.1653	0.4245	-0.6334	-.0319
(s-error)	0.187	0.169	0.254	0.106	0.237	0.092	0.066

Notes: The columns relate to the seven equations in the VECM, while the rows relate to the estimated adjustment coefficients. These coefficients correspond to the 3×7 matrix $(B_0)'$, where B_0 is defined in connection with equation (8) in the text. The variable $R(5)_{t-1}^\dagger$ contains lagged residuals from equation (10) in the text. Coefficients in bold are statistically significant at the 5% level of significance.

Table B3: Forecast error variance shares of real exchange rates (%)

Shocks	Horizons			
	h=0	h=3	h=6	h=9
Trade liberalization	9.21 [83.15]	11.61 [59.42]	12.26 [56.51]	13.48 [56.94]
Productivity improvement	0.15 [1.33]	0.74 [3.78]	0.98 [4.51]	1.01 [4.25]
Contractionary monetary policy	1.33 [12.01]	2.67 [13.65]	2.72 [12.53]	2.79 [11.79]
Expansionary government consumption	0.39 [3.51]	4.52 [23.15]	5.74 [26.45]	6.40 [27.02]
Total	11.08 [100.00]	19.54 [100.00]	21.70 [100.00]	23.67 [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 at the median of the posterior distribution.

Table B4: Comparison of the persistence of impulse responses produced by different approaches

Shocks	Approaches	<i>gov</i>	<i>open</i>	<i>x</i>	<i>y</i>	π	<i>si</i>	<i>q</i>
Trade liberalization	penalty function(1)	-(0-6)	+(0-9) [†]	+(0-1)	+(0-9)	-(0-2) [†]	-(0-1)	+(0-9)
	penalty function(2)	-(0-2)	+(0-9) [†]	-(0-9)	+(0-2)	-(0-2) [†]	-(0-1)	+(0-9)
	system-penalty fn.	No	+(0-9) [†]	No	No	-(0-1) [†]	No	No
	pure sign(4)	No	+(0-9) [†]	No	No	-(0-1) [†]	No	No
	pure sign(7)	-(0-2)	+(0-9) [†]	No	+(0-9)	-(0-1) [†]	-(0)	No
Productivity improvement	penalty function(1)	-(0-9)	No	+(0-9) [†]	+(0-9) [†]	-(0-1)	+(2-5)	-(2-4)
	penalty function(2)	-(0-9)	+(0-9)	+(0-9) [†]	+(0-9) [†]	-(0-1)	-(0)	No
	system-penalty fn.	No	No	+(0-9) [†]	+(0-9) [†]	-(0-1)	No	No
	pure sign(4)	No	No	+(0-3) [†]	+(0-9) [†]	No	No	No
	pure sign(7)	No	No	+(0-9) [†]	+(0-9) [†]	No	No	No
Contractionary monetary policy	penalty function(1)	No	-(0-1)	+(1-2) [†]	+(0-1) [†]	+(0-6) [†]	+(0-3) [†]	No
	penalty function(2)	No	-(0-1)	+(1-2) [†]	+(0-1) [†]	+(0-6) [†]	+(0-3) [†]	No
	system-penalty fn.	No	No	No [†]	+(0-9) [†]	+(0-6) [†]	+(0-2) [†]	No
	pure sign(4)	No	No	No [†]	+(0-9) [†]	+(0-5) [†]	+(0-2) [†]	No
	pure sign(7)	No	No	+(2-4) [†]	+(0-4) [†]	+(0-5) [†]	+(0-3) [†]	No
Expansionary government consumption	penalty function(1)	+(0-9) [†]	+(0-1)	-(0-9)	+(0-9) [†]	+(0-2) [†]	+(0-6)	-(0-9)
	penalty function(2)	+(0-9) [†]	+(0-1)	-(0-9)	+(0-9) [†]	+(0-2) [†]	+(0-6)	-(0-9)
	system-penalty fn.	+(0-9) [†]	No	No	+(0-6) [†]	+(0-2) [†]	No	No
	pure sign(4)	+(0-9) [†]	No	No	+(0-9) [†]	+(0-2) [†]	+(0-2)	No
	pure sign(7)	+(0-9) [†]	+(0-1)	No	+(0-2) [†]	+(0-2) [†]	No	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: (4) identify four shocks, (7) identify all seven shocks in the system.

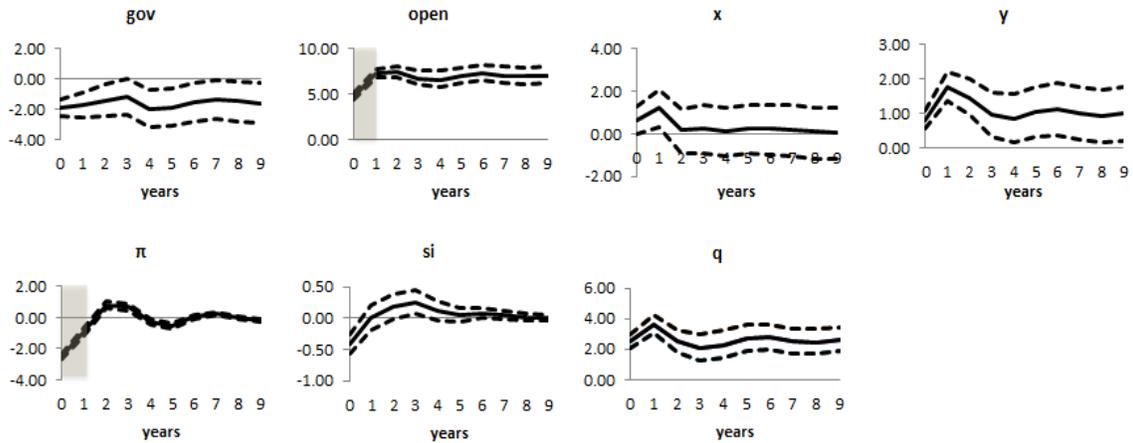
3) "[†]" indicates that the response is sign restricted

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

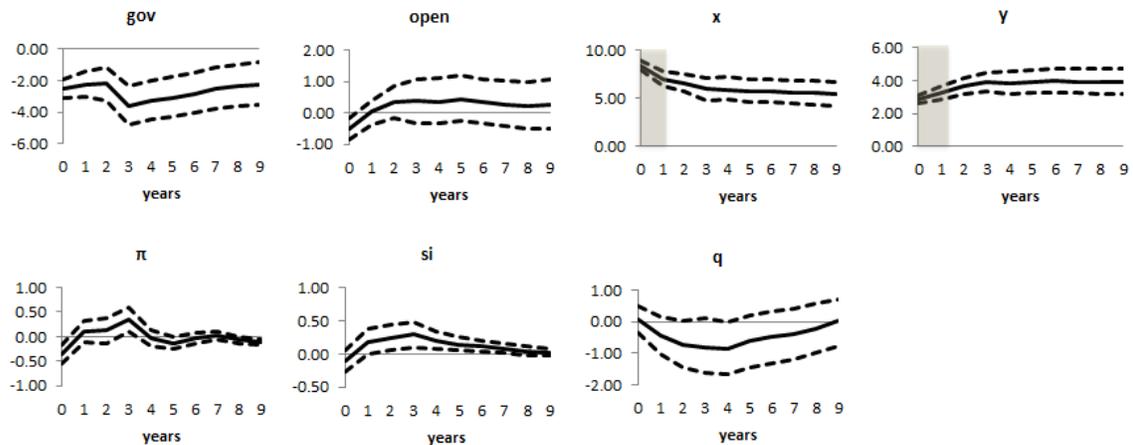
C Additional impulse response Analysis

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



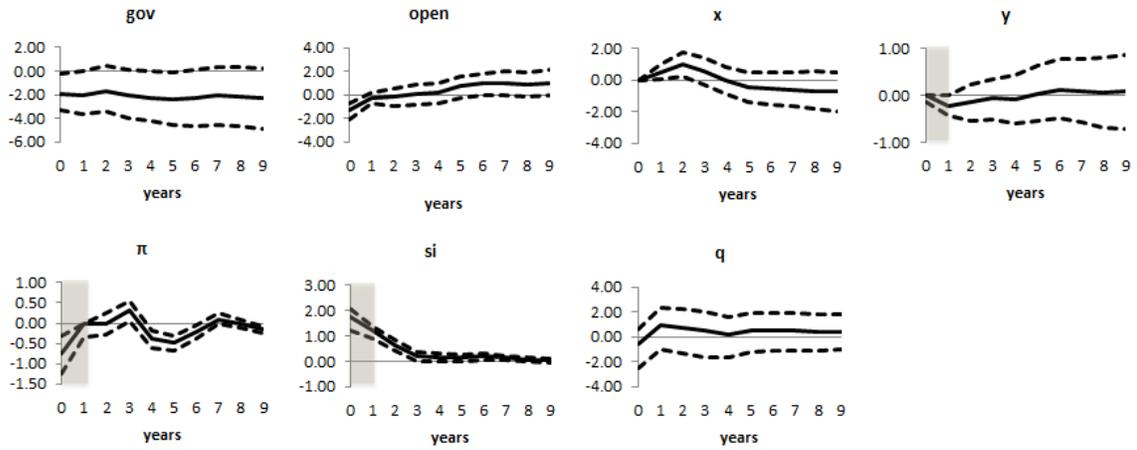
Notes: Each plot comprises of the solid line, which represents the median impulse response, and the dashed lines, which represent the 16% and 84% quantiles of the posterior distribution. The shaded areas indicate the responses restricted by sign restrictions.

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



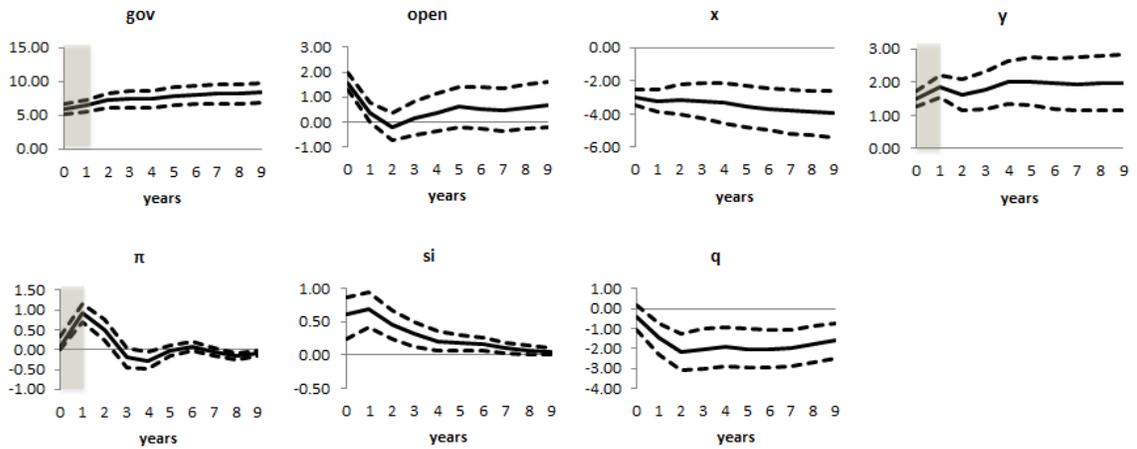
Notes: See Figure C1.

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.

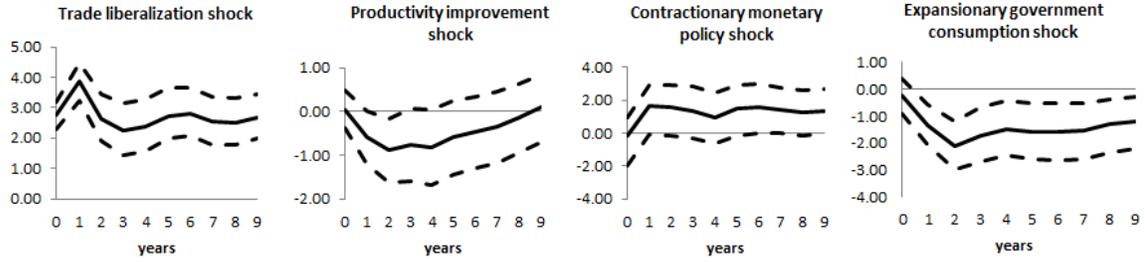
Figure C4: Impulse responses to an expansionary government consumption shock of one standard deviation in size, using sign restrictions with the penalty function.



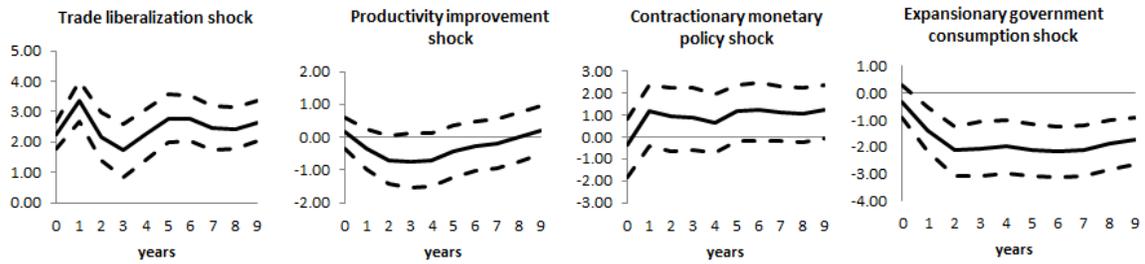
Notes: See Figure C1.

Figure C5: Robustness checks: Impulse responses of real exchange rates to the four shocks using various alterations.

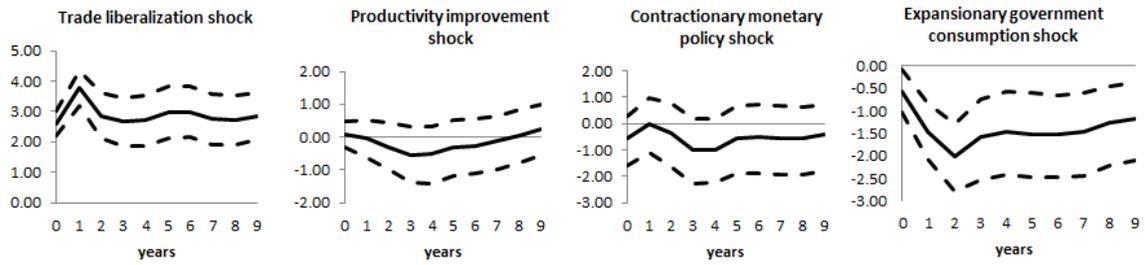
(1) Real exchange rate responses estimated from the model excluding China.



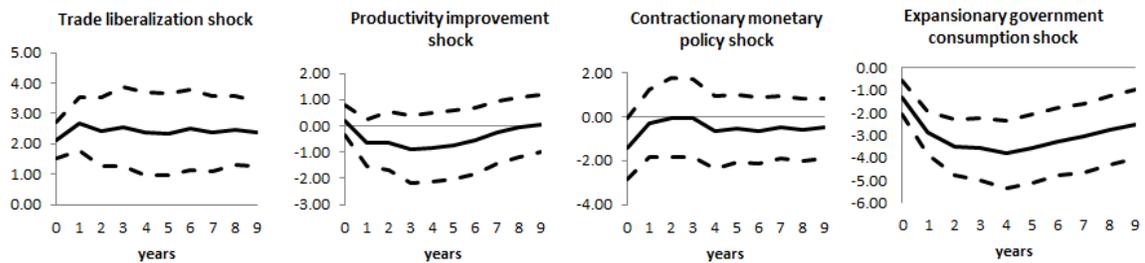
(2) Real exchange rate responses estimated from the model excluding India.



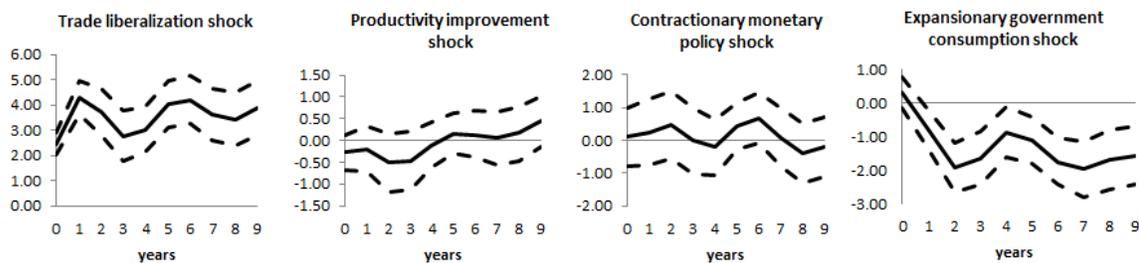
(3) Real exchange rate responses estimated from the model excluding Indonesia.



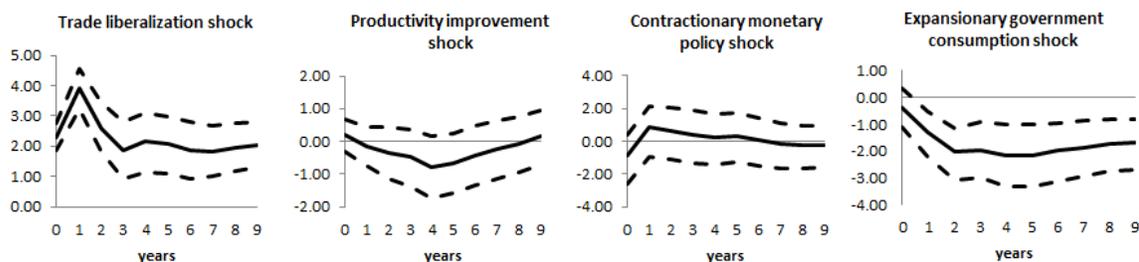
(4) Real exchange rate responses estimated by using the observations during the period 1970-1996.



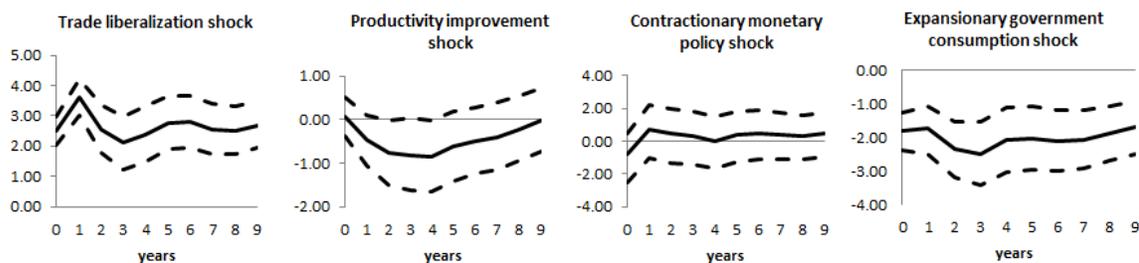
(5) Real exchange rate responses estimated by using the observations during the period 1997-2008.



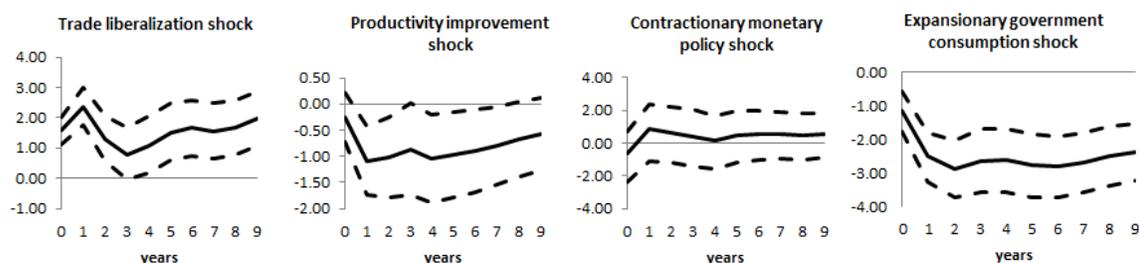
(6) Real exchange rate responses estimated from a VECMX*(4, 0).



(7) Real exchange rate responses based on identification scheme 1.



(8) Real exchange rate responses based on identification scheme 2.



(9) Real exchange rate responses estimated from a VARX* in levels.

