Real Exchange Rate Misalignments and U.S. Exports to Asia

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This paper examines real exchange rate misalignments in seven Asian countries and their impacts on U.S. exports. An analytical framework is first developed to estimate real exchange rate misalignments and, based on an autoregressive distributed lag (ARDL) bound testing approach, the long and short run relationships between U.S. exports and its key determinants are estimated. The empirical results indicate that real exchange rates are moderately misaligned in most of the Asian countries. A negative and significant long run relationship between real exchange rate misalignment and U.S. exports to each country is also detected. The results also show that real exchange rate continues to maintain its position as one of the important factors explaining international trade.

INTRODUCTION

One of the current challenges facing the world economy is the large and persistent trade imbalances believed to have been caused by competitive exchange rate policies adopted by various countries leading to currency misalignments. Empirical investigations show that more than a dozen of countries intervene in foreign exchange markets to deliberately devalue their currencies with the main objective of gaining competitive trade advantages. But given that an exchange rate is a shared variable, an attempt by one country to alter its value in a desired direction (say, a depreciation), creates the opposite effect (overvaluation) on the country's trading partners. Predatory currency market intervention is a form of modern-day mercantilist trade policy and has the same effect as import tariffs and export subsidy (Subramanian, 2008; Mattoo and Subramanian, 2009; Bergsten, 2007). As such, exchange rate manipulation, not only play an important role in a country's trade and overall economic performance, but also generate important international repercussions. The importance and severity of the current global imbalances require that adjustment takes place because, if left unchecked, they could pose a serious long-term problem for the global economy (Bank for International Settlement, 2006).

Contrary to the argument above that currency undervaluation can stimulate exports and retard imports in the undervalued-currency country and impose costs on trading partners, it can also be claimed that, under some conditions, currency undervaluation may have no perceptible effects on trading partners and, in fact, may positively contribute to their trade and welfare. Predatory currency manipulation to gain competitive advantage is a policy doomed to failure on both theoretical and empirical grounds and should have no effect on the manipulator and its trading partners alike. This proposition is reinforced in many academic debates that question whether a country can use real exchange rate as a policy variable to achieve a macroeconomic goal, and the fact that ascertaining the root cause of a currency misalignment is often a difficult matter in practice (Auboin and Ruta, 2011). According to Eichengreen (2008), monetary policy cannot be used to sustain a particular real exchange rate. Real exchange rates, according to the author, are dictated by economic fundamentals and cannot directly be controlled by policymakers. A controversy therefore arises about the trade and welfare effects of an undervalued currency. It then becomes crucial to determine whether an undervaluation policy by a country undermines or enhances the economic welfare of its trading partners (Staiger and Sykes, 2010).

The objective of this study is to contribute to the debate and, most importantly, quantify the effect of developing Asian countries' exchange rate policies on the exports of their trading partners (the United States in particular). The innovation in this research is that, contrary to previous studies that concentrate on evaluating the impacts of exchange rate movements on the country initiating the change, it focuses on the negative externalities a country's foreign exchange policy imposes on its trading partners in terms of lower exports. Given the long theoretical (and empirical, to a lesser extent) literature on the subject and the no less ambiguous evidence they produce, a reexamination of the relationship between exchange rate misalignments and the exports of trading partners is warranted.

The rest of the paper is structured as follows. A review of the literature on real exchange rate movements and international trade is presented in section 2. Section 3 provides the model and empirical results of the equilibrium RER and its misalignment. In section 4, the export model, econometric procedure, and export results are presented. Section 5 summarizes and wraps up the findings.

REAL EXCHANGE RATE MOVEMENTS AND INTERNATIONAL TRADE: A LITERATURE REVIEW

Real exchange rate (generally defined as $\frac{EP}{P^*}$, where P and P^* are the domestic and foreign price indexes, respectively, and E is the nominal exchange rate - units of foreign currency per domestic currency) is the relative price that matters for international trade. Understandably, simply altering the nominal exchange rate provides no guarantee that it will affect trade through its impact on the real exchange rate. The general belief is that exchange rate depreciation stimulates exports and retard imports. Movements in exchange rates may affect international trade by altering relative prices (a measure of real competitiveness), leading to expenditure switching by consumers and incentives for firms to reallocate resources between the tradable and non-tradable sectors of the economy. In that context, and assuming no market imperfections nor institutional weaknesses, exchange rate misalignments will have real effect on trade in the short run (as a result of the policy-induced temporary changes in relative prices), but should produce no real effect in the long run. This conventional view is scrutinized in recent theoretical and empirical works that raise serious skepticism and argue that the relationship between the level of a currency and trade is so complex and tedious that it warrants serious reflection. The effect of currency misalignment on international trade hinges critically on different factors, including the invoicing practices of the firms, the degree of price flexibility, how well market participants can anticipate exchange rate changes, the level of global production networks, and the extent of market imperfections (Auboin and Ruta, 2011; Staiger and Sykes, 2010; Bagwell, 1991; Freund and Pierola, 2010).

The currency in which domestic producers invoice their products is an important factor when assessing the trade effect of exchange rate movements. In the short run, although a devaluation may affect trade, the nature and magnitude of the effect depend on the invoicing practice (Staiger and Sykes, 2010). In the long-run, assuming no market imperfections, the currency in which the exporter invoices its products does not matter for international trade. Even the prediction that devaluation has no real effect in the long run is plagued with controversies as many distortions may prevent the adjustment of the relative prices in the long run, and that a number of specific empirical factors may cause exports to expand in the long run. Some studies provide evidence of positive impacts of exchange rate undervaluation on trade in the long run. One theory posits that currency undervaluation may induce firms to enter new foreign markets and produce new products, with long-run positive effects on exports (Freund and Pierola, 2010). Bernard and Jensen (2004) examined the sources of the manufacturing export boom for the U.S. during the 1987-1992 period and found that improvements in exchange rates and rising foreign incomes were the drivers of most of the export increase. Haddad and Pancaro (2010) on the other hand, found that real exchange rate undervaluation boosts exports only for countries with low per capita income, and only in the medium term. In the long run, however, the authors found that the effect of real undervaluation on exports is insignificant for all income levels.

The effect of exchange rate movements is also complicated by the channels through which it affects trade - its depreciation and variability (risk). Although exchange rate undervaluation may positively affect exports, the associated exchange rate risk it creates could offset or even dominate the positive effect, leading to a zero or negative net effect (Fang et al., 2006; Fang and Miller, 2004). Another important aspect that should not go unnoticed when analyzing the response of trade to changes in exchange rates, is the growth of cross-border production sharing whereby firms, in their production process, make use of large proportion of imported parts and components. The implication of this integrated production process is that the trade effect of a local currency undervaluation has two offsetting forces, with the net effect unknown and a matter of empirical investigation (Jongwanich 2009; Arndt and Huemer 2004; Abeysinghe and Yeok 1998; Athukorala and Suphachalasai 2004).

EQUILIBRIUM REAL EXCHANGE RATE AND MISALIGNMENT

Equilibrium RER Model and Cointegration Analysis

Given the complexity in estimating equilibrium real exchange rates and their misalignments, different theories and measures have been developed. The approach in this paper consists of (a) first estimating an equilibrium relationship between the RER and its economic fundamentals, (b) then using the estimated coefficients and the sustainable values of the fundamental variables to compute the equilibrium RER, and (c) finally, computing the exchange-rate misalignment as the deviation of the actual RER from the equilibrium RER.

Based on theoretical considerations, and following Elbadawi (1994), Kemme and Roy (2005), and Clark and MacDonald (1998), the equilibrium real exchange rate is hereby specified as a linear function of a set of fundamental economic variables as follows:

$$q^{*} = f\left(\underset{+}{G_N}, \underset{-}{G_T}, \underset{-}{NFA}, \underset{+/-}{INV}, \underset{-}{OPEN}, \underset{+}{PROD}, \underset{+/-}{TOT} \right) \tag{1}$$

where, q^{*} is the equilibrium real exchange rate, G_N and G_T are government spending on nontradable and tradable goods, respectively, NFA is net foreign assets, INV is investment, $OPEN$ stands for openness to trade, $PROD$ is relative productivity in the tradable goods sector relative to the nontradable sector, and TOT is terms of trade. The signs of the partial derivatives (expected signs of the coefficients) are shown below the variables in Equation (1). Because of data unavailability on government consumption on nontradable and tradable goods (G_N and G_T, respectively), government consumption ($GCON$) is used in place of the two. The openness to trade variable ($OPEN$) is proxied by the ratio of the sum of the real values of exports and imports to the value of gross domestic product.

As a prelude to estimating Equation (1), the variables are subject to different tests to ensure cointegration. An important precondition that must be satisfied prior to executing the cointegration test, is that all variables under consideration must be stationary and have the same order of integration, I(1). If stationarity is confirmed, Equation (1) is estimated using the Johansen (1988) cointegration test to determine the presence or absence of long-run relationship among the variables. The Johansen (1988) cointegration approach is based on two likelihood ratios (LR) test statistics - the trace statistic and the maximum eigenvalue statistic tests. The trace statistic tests the null hypothesis that “there are at most *r* cointegrating relations” against the alternative of “*n* cointegrating relations”. The maximum eigenvalue statistic tests the null hypothesis that “there are *r* cointegrating relations” against the alternative that “there are *r* +1 cointegrating relations”. The asymptotic critical values for both tests are tabulated in Johansen and Juselius (1990) and are also given by most econometric software packages.

Empirical Results of the Equilibrium RER and Misalignment

The empirical model is estimated for seven Asian countries (China, Hong Kong, India, Indonesia, Korea, Malaysia, and Thailand). One of the challenges in estimating the model for these Asian countries (as it is for most developing countries), is the unavailability, limited, or poor quality data. For some countries/series, data is available but only annually. In this case, the annual data is converted to quarterly frequency.

Nominal bilateral exchange rates, CPI, government spending, nominal GDP, investment, and real GDP are obtained from the IMF *International Financial Statistics*. GDP per capita and terms of trade data are derived from the *World Development Indicator* of the World Bank. Bilateral exports and imports are obtained from the *Direction of Trade Statistics* database of the International Monetary Fund. Bilateral real exchange rates are computed by the author.

To ensure stationarity as discussed above, the Augmented Dickey Fuller (ADF) unit root test is applied to each series to test for the existence of a unit root to determine the order of integration. The selection of the optimal lags in the test specification for each variable is obtained through a search process (i.e., the minimum Schwarz Information Criteria, SIC). Following Chen et al. (2008), all combinations with/without intercept and time trend are explored with the final choice determined by the parameter significance of each of the regressors. The results of the tests (not reported due to the size of the data) indicate clear evidence that all the variables are nontionationary in levels but stationary in first differences; that is, they are I(1).

The next step is to test for cointegration between the real exchange rate and its economic fundamentals, and determine the cointegration equation. The Johansen (1988) cointegration test is applied to Equation (1) for each country to determine the presence or absence of a long-run relationship. In order to select only models that best fit the data, different combinations of economic fundamentals are explored in the estimation of the relationship for each country, and the final choice based on the significance of the parameters and overall regression. Based on the results of the cointegration tests, the trace and maximum-eigenvalue statistics clearly point at the rejection of the null hypothesis of non-cointegrating vector at 5% level. The results therefore strongly confirm the existence of one cointegration vector for each country.
The normalized cointegration vectors are summarized in Table 1. For robustness check, the Lagrangian multiplier (LM) autocorrelation and the White heteroscedasticity analyses are also performed and the results shown in Table 1. The results of those two tests indicate that the model is structurally stable and there is no evidence of model misspecification for each of the seven countries. The cointegration results show that the long-run coefficients of all the fundamental variables are statistically significant and correctly signed (except for the $GCON$ coefficient which is negative for China, Malaysia, and Japan). These results confirm the theory and previous empirical studies that, in the long run, real exchange rates are driven by economic fundamentals.

The equilibrium real exchange rate is obtained by feeding the estimated model (Equation 1) with the permanent components of the fundamentals. The equilibrium RER is therefore expressed as follows:

$$q^{*} = \beta^{'} F^{p} \tag{2}$$

where q^{*} is the equilibrium RER, β is the vector of the estimated coefficients from Equation (1), and F^{P} is the permanent or sustained values for the fundamentals. The Hodrick-Prescott (1997) smoothing technique (H-P filter) is used in this study to remove the short time fluctuations of the economic fundamentals, thereby revealing the long-run time trends of the series.

TABLE 1
COINTEGRATION REGRESSION RESULTS

China
RER = -1.606 -0.844*GCON* -0.330*OPEN* -0.137*NFA* +0.179*PROD* -0.802*TOT*
(-6.92)* (-8.45)* (-1.97)** (3.79)* (-11.03)*
LM(1):p-value = 0.71 LM(2):p-value = 0.74 White Heteroscedasticity: p-value = 0.97
Sample Period: 1986Q1-2013Q4

Hong Kong
RER = -9.947 +0.946*GCON* +1.416*INV* +2.587*PROD* -0.015*TOT* -0.004*TREND*
(6.25)* (9.62)* (6.86)* (-4.77)* (-4.27)*
LM(1):p-value = 0.82 LM(2):p-value = 0.28 White Heteroscedasticity: p-value = 0.08
Sample Period: 1985Q1-2013Q4

India
RER = -4.788 +1.353*GCON* -1.487*OPEN* +1.302*PROD* -1.296*TOT*
(7.68)* (-6.66)* (4.24)* (-7.58)*
LM(1):p-value = 0.07 LM(2):p-value = 0.14 White Heteroscedasticity: p-value = 0.10
Sample Period: 1981Q1-2013Q4

Indonesia
RER = -12.882 +1.565*GCON* -1.122*OPEN* -0.393*INV* +3.423*PROD* +2.752*TOT* -0.010*TREND*
(7.02)* (-4.85)* (-2.11)** (8.41)* (5.39)* (-8.60)*
LM(1):p-value = 0.12 LM(2):p-value = 0.57 White Heteroscedasticity: p-value = 0.57
Sample Period: 1990Q1-2013Q4

Korea
RER = -3.948 +3.533*GCON* -1.854*OPEN* -6.659*INV* +0.297*PROD* -6.673*TOT*
(1.57) (-1.81)*** (-4.31)* (4.83)* (-2.34)**
LM(1):p-value = 0.28 LM(2):p-value = 0.91 White Heteroscedasticity: p-value = 0.32
Sample Period: 1981Q1-2013Q4

Malaysia
RER = -7.555 -0.178*GCON* -0.307*OPEN* +0.166*INV* +2.031*PROD* +3.604*TOT* -0.008*TREND*
(-2.62)* (-3.17)* (3.12)* (8.04)* (8.04)* (-7.52)*
LM(1):p-value = 0.05 LM(2):p-value = 0.12 White Heteroscedasticity: p-value = 0.39
Sample Period: 1991Q1-2013Q4

Thailand
RER = -7.227 +1.042*GCON* -2.062*OPEN* +1.918*PROD* +0.134*TOT*
(2.50)** (-7.01)* (3.91)* (2.22)**
LM(1):p-value = 0.37 LM(2):p-value = 0.10 White Heteroscedasticity: p-value = 0.05
Sample Period: 1993Q1-2013Q4

Notes:
RER: An increase is an appreciation of the domestic (Asian) currency.
Numbers in parentheses below the estimated coefficients are t-statistics.
*, **, *** denote statistical significance at 1%, 5%, and 10% levels, respectively.

The degree of misalignment ($RMIS$) is then derived as the deviations of the actual (observed) RER from its equilibrium level:

$$RMIS = q - q^*$$

where q and q^* are the actual and equilibrium RERs, respectively.

Given the definitions of the RER and misalignment ($RMIS$) variables as expressed above, when $RMIS$ is positive, it represents a real exchange rate overvaluation of the currency of the Asian country in question (a depreciation of the U.S. dollar). In contrast, there is a real exchange rate undervaluation of the Asian currency when $RMIS$ is negative. The estimated RER misalignment ($RMIS$) series are used in the next section to determine their impacts on U.S. exports to each of the seven Asian countries.

RER MISALIGNMENTS AND U.S. EXPORTS PERFORMANCE

Export Model and Econometric Procedure

Based on traditional international trade theory, it is postulated that export demand is a function of domestic and foreign real incomes, and the level of real exchange rate. To address the issue at hand, this standard export demand-based framework is extended by including real exchange rate misalignment as determinant of real export, consistent with previous empirical studies, (Sapir and Sekkat, 1995; Mckenzie, 1998; Sekkat and Varoudakis, 2000; Doraisami, 2004; Kumakura, 2005; Bryne et al., 2008). The reduced-form export equation is therefore expressed a follows:

$$X_t^A = \alpha_0 + \alpha_1 \underset{+}{RER_t} + \alpha_2 \underset{+}{Y_t^A} + \alpha_3 \underset{+}{Y_t^{US}} + \alpha_4 \underset{-}{RMIS_t} + \xi_t \qquad (3)$$

where,

X^A = U.S. export volume to Asian country A

RER = Real exchange rate

Y^A = Real income in Asian country A

Y^{US} = Real income in U.S. (A variable that captures U.S. capacity utilization and demand pressure on U.S. export performance)

$RMIS$ = Real exchange rate misalignment

ξ = White noise error term

To estimate Equation (3), it is crucial that we first test and establish cointegration amongst the variables included in the model. The Johansen test is widely used to test for cointegration. However, an important precondition that must be satisfied prior to executing the Johansen cointegration test is that all the variables under consideration must have the same order of integration, that is I(1). The time series properties of the variables included in Equation (3) are therefore first investigated. The stationarity test however reveals that the variables do not have the same order of integration. All but the RER misalignment variable (RMIS) are integrated of order one, I(1). The RER misalignment variable for each of the seven Asian countries under investigation, on the other hand, is stationary; that is, I(0).

Given the different orders of integration for the variables in the export equation, the Johansen cointegration test is not applicable. To establish long run relationship, this study relies instead on the autoregressive distributed lag (ARDL) bounds testing approach developed by Pesaran and Shin (1999) and Pesaran et al (2001). The bounds test can be applied to any set of data, whether purely I(0), purely I(1), or a mixture of the two. The procedure allows different variables to be assigned different lag-lengths as they enter the model, and is more robust even in small sample cases (Pesaran et al., 2001; Narayan, 2005).

To implement the bounds testing approach to investigate the long run relationship amongst the variables, the export Equation (3) is modelled as a conditional or unrestricted ARDL and takes the following form:

$$\Delta X_t = a_0 + \delta_1 X_{t-1} + \delta_2 RER_{t-1} + \delta_3 Y_{t-1}^{A} + \delta_4 Y_{t-1}^{US} + \delta_5 RMIS_{t-1} + \sum_{i=1}^{n} \theta_1 \Delta X_{t-i}$$
$$+ \sum_{i=0}^{n} \theta_2 \Delta RER_{t-i} + \sum_{i=0}^{n} \theta_3 \Delta Y_{t-i}^{A} + \sum_{i=0}^{n} \theta_4 \Delta Y_{t-i}^{US} + \sum_{i=0}^{n} \theta_5 \Delta RMIS_{t-i} + \varepsilon_t \tag{4}$$

where a_0 is the drift term, δ_j (j =1, 2, 3, 4, 5) are the long run coefficients, and ε_t is the white noise error term, and n is the lag length. The lag length can be determined based on the Akaike Information Criteria (AIC) or Schwarz Information Criteria (SIC) through the process of setting a priori a maximum lag length, with the optimal lag length being the one with the lowest AIC or SIC value. This procedure may introduce a potential bias since it requires setting the maximum lag order a priori while the true lag order of the model is unknown a priori (Pesaran, 2001). Following previous studies (Sidek et al., 2010; Wong and Tang, 2008), and to avoid such potential bias, four or five lag lengths are used in this study, depending on the one that best fits the data.

Cointegration amongst the variables in Equation (3) is tested using the Wald test. This involves testing the null hypothesis of cointegration against the alternative hypothesis, by calculating an F-statistics for the joint significance of the estimated long-run coefficients of the one-period lagged level of the variables in Equation (4). The calculated F-statistic is compared to the critical values tabulated by Pesaran et al. (2001). If the computed F-statistic falls below the lower critical bound, then the null hypothesis of no cointegration cannot be rejected. If the computed F-statistic is greater than the upper critical bound, then the null hypothesis can be rejected, suggesting cointegration (or that a long-run relationship exists amongst the variables of interest). If the obtained F-statistic falls within the two critical bounds, then the cointegration test is inconclusive.

Assuming that, as a result of the bounds testing approach to cointegration, a long-run relationship is detected amongst the variables, the next step is to estimate the following long-run export function to get the long-run coefficients, λ_j (j = 1, 2, 3, 4, 5).

$$\Delta X_t = b_0 + \sum_{i=1}^{n} \lambda_1 \Delta X_{t-i} + \sum_{i=0}^{n} \lambda_2 \Delta RER_{t-i} + \sum_{i=0}^{n} \lambda_3 \Delta Y_{t-i}^{A}$$
$$+ \sum_{i=0}^{n} \lambda_4 \Delta Y_{t-i}^{US} + \sum_{i=0}^{n} \lambda_5 \Delta RMIS_{t-i} + \mu_t \tag{5}$$

where b_0 is a constant term, and μ_t is the white noise error term.

In the final step, the dynamic short-run relationship between export volume and its key determinants is obtained using the following error correction model of the ARDL approach.

$$\Delta X_t = c_0 + \psi ECT_{t-1} + \sum_{i=1}^{n} \tau_1 \Delta X_{t-i} + \sum_{i=0}^{n} \tau_2 \Delta RER_{t-i} + \sum_{i=0}^{n} \tau_3 \Delta Y_{t-i}^{A}$$
$$+ \sum_{i=0}^{n} \tau_4 \Delta Y_{t-i}^{US} + \sum_{i=0}^{n} \tau_5 \Delta RMIS_{t-i} + \zeta_t \tag{6}$$

where c_0 is a constant term, ECT_{t-1} is the error correction term, ψ is a measure of the speed of adjustment to the long-run equilibrium, τ_j (j = 1, 2, 3, 4, 5) are the short-run dynamic coefficients, and ζ_t is the white noise error term.

As in the long-run model above, a maximum lag length of four or five is initially imposed on Equation (6). The parsimonious ARDL was selected based on the general-to-specific model (GSM) selection strategy. The model requires progressively eliminating insignificant lags from the equation, while checking for misspecification at each level (Hendry et. al., 1984; Wickens and Breusch, 1988; Hendry, 1995; Pesaran et al. 2001). The main advantage of this modelling strategy is that the final model is supported by the data and is free of "omitted variable" bias (see Gilbert, 1986).

Empirical Results of the Export Equations

Quarterly data are used to estimate the export equations for the seven Asian countries during different time periods. The export volume variable is obtained by deflating U.S. bilateral nominal export value by the U.S. export price index. The bilateral nominal export values are gathered from the IMF *Direction of Trade Statistics*. The U.S. export price index, U.S. real income and real income for Asian countries are derived from the IMF *International Financial Statistics*. For countries that do not have quarterly data, yearly data are used to derive quarterly series.

Prior to estimating the long and short run export equations, the ARDL bounds test is first considered to check for cointegration amongst the variables of interest for each country. The bounds test results are reported in Table 2.

TABLE 2
BOUNDS TEST FOR COINTEGRATION: F-STATISTICS AND CRITICAL VALUES

	China	**Hong Kong**	**India**	**Indonesia**	**Korea**	**Malaysia**	**Thailand**
F-Statistic	3.95^{***}	5.13^{*}	5.17^{*}	4.74^{**}	4.76^{**}	5.04^{**}	3.94^{***}
Critical Values†		*Lower Bound, I(0)*	*Upper Bound, I(1)*				
	1%	3.74	5.06				
	5%	2.86	4.01				
	10%	2.45	3.52				

Notes:
†: Critical values based on k = 4 explanatory Variables.
*, ** and *** denote statistical significance at 1%, 5%, and 10% levels, respectively.

As the results indicate, the calculated *F-statistics* are greater than the upper bound of the critical values, indicating that the variables are cointegrated and, hence, there is a long-run relationship between export and its key determinants for each of the seven Asian countries. The results are significant at the 1% level (Hong Kong and India), 5% level (India, Korea, and Malaysia), and 10% level (China and Thailand).

Confident about the existence of long run relationship amongst the variables, the long run coefficients are estimated. The empirical results of the cointegration equation for each country are presented in Table 3. Most of the variables have the expected signs and are statistically significant. A positive and statistically significant (at 1% and 5% levels) relationship between real exchange rate and exports is revealed for each of the countries. Consistent with previous studies, this result strongly reveals that a real exchange rate depreciation encourages exports in the devaluing country (Asian country in the case of this study), but most importantly, represents a serious hindrance for exports in trading-partner countries (United States in this study). The empirical results show that the range of the negative impact on U.S.

exports of a 1% RER depreciation in an Asian country is 0.31% (Korea) to as much as 4.76% (Malaysia). The impact on U.S. exports (as expressed by the RER coefficients) obviously varies depending on the country where the depreciation occurs. Low impact (that is, less than 1% reduction in U.S. exports for each 1% RER depreciation of an Asian currency) is revealed for China, Indonesia, and Korea, and high impact (more than 1% reduction in U.S. exports for each 1% RER depreciation in an Asian currency) countries are represented by Hong Kong, India, Malaysia, and Thailand.

TABLE 3
U.S. EXPORTS PERFORMANCE: ESTIMATED LONG RUN RELATIONSHIPS

	China	Hong Kong	India	Indonesia	Korea	Malaysia	Thailand
Constant	-2.020	-2.212	3.445	0.843	0.470	0.221	0.069
	(-3.30)	(-5.77)	(3.31)	(0.73)	(0.38)	(0.13)	(0.05)
RER	0.498**	1.519**	3.929*	0.517**	0.305***	4.762*	4.333*
	(2.01)	(2.40)	(8.00)	(2.46)	(1.80)	(5.09)	(3.38)
Y^A	0.730*	0.418	0.422*	1.257*	0.558*	-4.098*	-6.538*
	(3.82)	(0.90)	(3.50)	(3.51)	(10.94)	(-3.80)	(-3.18)
Y^{US}	0.007***	0.930**	0.875*	-0.001	0.097	4.273*	3.390*
	(1.93)	(2.51)	(7.93)	(0.39)	(0.43)	(3.68)	(2.79)
RMIS	-0.100*	-0.027***	-0.032***	-0.014**	-0.008***	-0.017**	-0.354*
	(-7.10)	(-1.69)	(-1.87)	(-2.08)	(-1.90)	(-2.20)	(-3.51)

Notes:
RER: An increase signifies an appreciation of Asian currency.
A decrease represents a depreciation of Asian currency.
Numbers in parentheses are t-statistics.
*, ** and *** denote statistical significance at 1%, 5%, and 10% levels, respectively.

As pointed out above, there are theories and empirical studies that explain the weak relationship (small but not necessarily statistically insignificant RER coefficient) between RER and exports, as is the case for China, Indonesia, and Korea in this study. Arndt and Huemer (2007) posits that for international trades where goods flow within production networks, the sensitivity of trade to RER is altered. According to the authors, the effect of RER on U.S. exports of goods with high proportion of parts and components (to be assembled into products that are then exported rather than used at home) should therefore be weaker. In their empirical investigation of the Thai export performance, Athukorala and Suphachalasai (2004) found a relatively low elasticity of machinery and transports equipment exports with respect to real exchange rate, consistent with their observation of an ongoing process of product fragmentation in the export expansion of that product category. Jongwanich (2009) reached similar conclusion for Singapore whose RER has a statistically insignificant effect on the country's exports, a reflection of the high proportion of manufacturing parts and components in the country's exports and imports. In the case of China, the relatively low coefficient obtained in this study for the RER may be explained, at least partially, by the production networks phenomenon that progressively and consistently describes the international trade environment between the United States and China.

Regarding the RER misalignment variable, it is expected to affect real exports in two ways. The first channel is through the price effect that has a direct and immediate impact on exports. A real overvaluation signifies a loss in international competitiveness that can harm the country's exports. A persistent real undervaluation may also hurt a country's exports as the undervaluation is expected to lead to an overheated economy, higher imports prices and thus domestic prices, and eventually, an appreciation of the currency. The second channel is the more indirect but important channel when the RER misalignment causes misallocations of resources from the tradable to the nontradable sectors of the economy.

A negative and statistically significant relationship is found in the long run between RER misalignments ($RMIS$) and U.S. real exports to each of the seven Asian countries. This clear verdict strongly indicates that RER misalignments of Asian currencies have adverse effects on U.S. real exports to those countries. The magnitude of the impact in the long run is however relatively low to moderate. The impact on U.S. exports for every 1% RER misalignment, ranges from 0.008% (Korea) to 0.35% (Thailand). The total adverse impact of real exchange rate on U.S. exports could however be severe if the Asian currencies are misaligned during a time of persistent real undervaluation of those currencies. The results show that a 1% real exchange rate misalignment of the Chinese Renminbi will lower U.S. real exports to China by 0.10% in addition to a 0.50% reduction in U.S. exports if the real misalignment is coupled with a 1% real depreciation of the yuan. On the other hand, those numbers would be 0.35% and 4.33%, respectively, for Thailand.

The coefficients of the last two determinants of real exports in the long run are mostly significant but with mixed signs. Given the ongoing production networks and product fragmentation, domestic and foreign incomes are likely to explain a country's exports. The results in this study largely confirm the increasing role played by domestic income in explaining exports. The coefficient of the U.S. income is positive in six of the seven Asian countries and statistically significant in five. The results underscore the increasing role played by domestic income in explaining not only imports but also a country's exports. The results also show that an increase in real incomes in Asian countries boots U.S. exports to those countries, except for Malaysian and Thailand. Positive and statistical significant coefficient of foreign income variable is expected according to the traditional trade model. The negative and statistically significant coefficients for the Malaysia and Thailand real incomes (Y^A) is inconsistent with traditional export models. Negative foreign income coefficients were also found in McKenzie and Brooks (1997) and Arndt and Huemer (2007). According to Arndt and Huemer (2007), in the context of cross-border production sharing, for a given country, variations in foreign income should be unimportant in explaining the country's exports, while the country's GDP should play a larger role in explaining its exports. So, the negative sign of the foreign income (Malaysia and Thailand) in this study is not the surprise but its statistical significance.

Moving on to the dynamic short run estimates, the results are reported in Table 4. The reliability of these error correction models are documented through a number of diagnostic tests including, Breusch-Godfrey LM test for serial correlation, Jarque-Bera (JB) test for normality, Ramsey RESET test for misspecification, and the ARCH test for heteroscedasticity.

The results of these diagnostic tests, also presented in Table 4, show that the short run dynamic models are well specified, suggesting that the models do not suffer from serial correlation, heteroscedasticity, misspecification error, and that they are stable at reasonable significance levels.

The coefficient of the error-correction term (ECT) is negative as expected across all models, and statistically significant at 1% level for all countries, except Thailand where it is significant at 10% level. These negative signs and statistical significance are further evidence of cointegration. The magnitude of the coefficient of the error correction term, a measure of the speed of adjustment of short run fluctuations toward long run equilibrium, ranges from 0.07 (Thailand) to 0.95 (Indonesia). These results show that, except for Thailand, the speed of adjustment to long run equilibrium is relatively fast.

Consistent with the long run estimates, positive RER coefficients are detected in the short run for India and Korea. For the remaining countries, the signs of the RER coefficients are mixed. Except for Hong Kong, the magnitude of the short run RER coefficients are similar in magnitude compared to the long run coefficients. Similar results are obtained for the RER misalignment variable ($RMIS$). Negative signs are detected for India, Indonesia, and Thailand, positive signs for China, and a mixture for the remaining countries. Perhaps most importantly, is the similarity between the magnitudes (low to moderate) of the RER misalignment coefficients in the long and short run.

TABLE 4
U.S. EXPORTS PERFORMANCE: ESTIMATED SHORT RUN ERROR CORRECTION MODEL

China

$$\Delta X = 0.04 - 0.33ECM_{t-1} - 0.33\Delta X_{t-1} - 0.22\Delta X_{t-2} + 0.22\Delta X_{t-4} + 0.52\Delta RER_{t-3} - 0.56\Delta RER_{t-4} + 8.30Y^{A}_{t-1}$$

(1.66) (-4.05)* (-3.63)* (-2.69)* (3.03)* (2.09)** (-2.21)** (-2.16)**

$$- 9.36\Delta Y^{A}_{t-2} + 18.44Y^{A}_{t-3} - 10.53Y^{A}_{t-4} - 0.24Y^{US}_{t-4} + 0.02\Delta RMIS_{t-1} + 0.02\Delta RMIS_{t-2}$$

(1.81)*** (3.54)* (-2.50)** (-2.45)** (2.58)* (2.46)**

$$+ 0.02\Delta RMIS_{t-3} + 0.01\Delta RMIS_{t-4}$$

(3.08)* (1.83)***

$R^2 = 0.58$ $\overline{R^2} = 0.52$ LM(1)=0.64(0.43) LM(2)=0.39(0.68) ARCH=1.47(0.22) JB=2.64(0.27)
RESET=1.33(0.27)

Hong Kong

$$\Delta X = -0.01 - 0.66ECM_{t-1} - 0.25\Delta X_{t-4} + 29.27\Delta RER_{t-1} - 32.51\Delta RER_{t-2} + 49.26\Delta RER_{t-3} - 24.69\Delta RER_{t-4}$$

(-0.90) (-7.26)* (-2.87)* (3.32)* (-2.65)* (3.93)* (-2.42)**

$$- 3.35\Delta RER_{t-5} + 0.62Y^{A} - 0.67Y^{A}_{t-1} + 1.10Y^{US} + 0.92Y^{US}_{t-1} - 1.01Y^{US}_{t-2} - 0.04\Delta RMIS + 0.25\Delta RMIS_{t-1}$$

(-3.99)* (3.64)* (-3.74)* (2.03)** (1.61) (-1.86)*** (-5.19)* (3.27)*

$$- 0.26\Delta RMIS_{t-2} + 0.43\Delta RMIS_{t-3} - 0.19\Delta RMIS_{t-4}$$

(-2.49)** (4.02)* (-2.13)**

$R^2 = 0.74$ $\overline{R^2} = 0.67$ LM(1)=0.34(0.56) LM(2)=0.31(0.74) ARCH=0.09(0.91) JB=3.03(0.22)
RESET=0.08(0.78)

India

$$\Delta X = -0.01 - 0.45ECM_{t-1} - 0.30\Delta X_{t-1} - 0.24\Delta X_{t-2} + 1.12\Delta RER + 0.48\Delta RER_{t-3} - 0.21\Delta Y^{A}_{t-1} - 0.45\Delta Y^{A}_{t-2}$$

(-1.15) (-5.76)* (-3.00)* (-2.70)* (2.43)** (1.14) (-2.62)* (-5.62)*

$$- 0.24Y^{A}_{t-3} - 2.34\Delta Y^{US}_{t-1} - 2.98\Delta Y^{US}_{t-2} - 2.20\Delta Y^{US}_{t-3} - 0.01\Delta RMIS$$

(-2.73)* (-2.06)** (-2.55)** (-1.95)*** (-1.68)***

$R^2 = 0.57$ $\overline{R^2} = 0.50$ LM(1)=0.06(0.81) LM(2)=0.74(0.48) ARCH=0.93(0.43) JB=0.44(0.80)
RESET=1.03(0.31)

Indonesia

$$\Delta X = -0.05 - 0.95ECM_{t-1} + 0.14\Delta X_{t-1} + 0.26\Delta X_{t-2} + 0.24\Delta X_{t-4} + 0.22\Delta X_{t-5} - 0.32\Delta RER_{t-2} - 7.30\Delta Y^{A}_{t-3}$$

(-1.37) (-6.78)* (1.14) (2.38)** (2.71)* (2.27)** (-1.50) (-2.52)**

$$+ 10.96\Delta Y^{A}_{t-4} - 4.83\Delta Y^{A}_{t-5} + 3.4\Delta Y^{US} - 7.64^{US}_{t-2} + 4.72\Delta Y^{US}_{t-3} - 0.03\Delta RMIS$$

(3.09)* (-1.81)*** (2.64)* (-2.31)** (2.14)** (-2.94)*

$R^2 = 0.60$ $\overline{R^2} = 0.53$ LM(1)=0.06(0.80) LM(2)0.18=(0.83) ARCH=0.76(0.56) JB=2.60(0.27)
RESET=0.04(0.84)

TABLE 4 (Cont.)
U.S. EXPORTS PERFORMANCE: ESTIMATED SHORT RUN ERROR CORRECTION MODEL

Korea

$$\Delta X = -0.01 - 0.75ECM_{t-1} - 0.14\Delta X_{t-1} + 0.08\Delta X_{t-3} + 0.73\Delta RER + 0.51\Delta RER_{t-2} + 1.98\Delta RER_{t-4} + 0.24\Delta Y^{A}_{t-1}$$
(-1.37) (-6.85)* (-1.34) (1.05) (5.81)* (3.88)* (4.85)* (1.31)

$$+ 0.41\Delta Y^{A}_{t-2} + 0.34\Delta Y^{A}_{t-3} + 0.79\Delta Y^{A}_{t-4} + 0.34\Delta Y^{US}_{t-1} - 0.36\Delta Y^{US}_{t-4} - 0.02\Delta RMIS_{t-1} + 0.07\Delta RMIS_{t-4}$$
(2.36)** (1.79)*** (4.28)* (1.80)*** (-1.86)*** (-3.23) * (5.27)*

$R^2 = 0.77$ $\overline{R}^2 = 0.72$ LM(1)=0.76(0.37) LM(2)=0.38(0.69) ARCH=0.11(0.90) JB=0.84(0.66)*
RESET=0.17(0.68)

Malaysia

$$\Delta X = -0.02 - 0.20ECM_{t-1} - 0.20\Delta X_{t-2} - 0.20\Delta X_{t-3} - 0.21\Delta X_{t-5} + 1.15\Delta RER + 1.84\Delta RER_{t-1} - 1.09\Delta RER_{t-3}$$
(-1.51) (-3.10)* (-1.71)*** (-1.85)*** (-1.96)** (1.75)*** (2.69)* (-1.78)***

$$-1.00\Delta RER_{t-4} + 1.60\Delta RER_{t-5} - 5.03\Delta Y^{A} + 8.26\Delta Y^{A}_{t-1} - 3.09\Delta Y^{A}_{t-2} - 2.93\Delta Y^{A}_{t-4} + 2.54\Delta Y^{A}_{t-5}$$
(-1.96)** (4.14)* (-3.39)* (4.76)* (-2.48)** (-2.03)** (1.97)**

$$+ 0.01\Delta RMIS_{t-1} + 0.01\Delta RMIS_{t-2} - 0.01\Delta RMIS_{t-4} + 0.02\Delta RMIS_{t-5}$$
(2.27)** (2.16)** (-1.00) (2.87)*

$R^2 = 0.62$ $\overline{R}^2 = 0.45$ LM(1)=1.38(0.25) LM(2)=0.68(0.51) ARCH=0.32(0.73) JB=0.31(0.85)
RESET=0.03(0.85)

Thailand

$$\Delta X = -0.02 - 0.07ECM_{t-1} - 0.54\Delta X_{t-1} - 0.33\Delta X_{t-2} - 0.27\Delta X_{t-3} - 0.50\Delta RER + 0.94\Delta RER_{t-2} + 0.45\Delta RER_{t-4}$$
(-1.39) (-1.81)*** (-4.48)* (-2.53)** (-2.27)** (-1.58) (2.65)* (1.07)

$$+1.40\Delta Y^{A} + 2.12\Delta Y^{A}_{t-1} + 1.49\Delta Y^{A}_{t-2} + 1.56\Delta Y^{A}_{t-3} + 1.46\Delta Y^{A}_{t-4} + 2.51\Delta Y^{US} + 1.00\Delta Y^{US}_{t-3} - 0.03\Delta RMIS$$
(2.85*) (3.37)* (2.38)** (2.87)* (2.84)* (2.45)** (0.92) (-1.79)***

$$-0.03\Delta RMIS_{t-2} - 0.02\Delta RMIS_{t-4}$$
(-2.15)** (-1.50)

$R^2 = 0.62$ $\overline{R}^2 = 0.51$ LM(1)=0.62(0.43) LM(2)=0.36(0.70) ARCH=0.59(0.0.56) JB=2.30(0.32)
RESET=0.84(0.36)

Notes: *, ** and *** denote statistical significance at 1%, 5%, and 10% levels, respectively.

CONCLUSIONS

This study examines the impact of real exchange rate misalignments in seven Asian countries (namely, China, Hong Kong, India, Indonesia, Korea, Malaysia, and Thailand) on the exports of their trading partner, the United States. To that end, an analytical framework is first developed to estimate the real exchange rate misalignment for each of the seven countries. To establish the long run relationship between U.S. exports and its key determinants and estimate the cointegration vectors, the study relies on the autoregressive distributed lag (ARDL) bounds testing approach. In the final step, the dynamic short-run relationship is obtained using the error correction model of the ARDL approach.

The results indicate that real exchange rates are misaligned in most of the Asian countries, though not to the extent claimed in some studies. A negative and significant long run relationship between real exchange rate misalignment and U.S. exports to each country is also detected and that relationship is robust across all models. The results also show that real exchange rate continues to maintain its position as one of the most important factors that explain international trade. Consistent with previous studies, a positive and statistically significant relationship between real exchange rate and exports is revealed for each of the seven Asian countries, suggesting that predatory currency manipulation by Asian countries, if materialized, could cost the United States in terms of exports and jobs. The total adverse impact on U.S. exports could be worse if the Asian currencies are misaligned during a time of persistent real undervaluation of those currencies.

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