The Impact of a Common Currency on East Asian Production Networks and China's Exports Behavior [*]

By Mizanur Rahman[‡]

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Abstract:

This paper estimates the impact of a common currency on East Asian exports, particularly by focusing on China's processing exports. The rapid growth in these exports clearly indicates China's deepening integration with other East Asian countries and also to the rest of the world. The results show that the cost to China's processing exports for not having a common currency is more than double the corresponding cost to China's ordinary exports. The magnitudes of these costs are consistent with the hypothesis that a common currency in East Asia would further integrate East Asian production networks and promote those exports whose value chains are spatially fragmented across borders in East Asia.

Keywords: East Asian production networks, fragmentation of value chain, currency area, common currency, exchange rate management, China.
JEL classification: F14, F33, F36, F42

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[‡] Assistant Professor, University of Dhaka, Bangladesh; PhD Scholar, National Graduate Institute for Policy Studies (GRIPS), 7-22-1 Roppongi, Minato-ku, Tokyo 106-8677; and Researcher, Research Institute for Economy, Trade, and Industry (RIETI), Japan. Tel: +81(3)-3501-0230; Fax: +81(3)-3501-8577; E-mail: D0305@grips.ac.jp <or> mizanur-rahman@rieti.go.jp

1. Introduction

In the context of globalization, the surging of regional trade agreements (RTAs) and preferential trade agreements (PTAs) has raised several questions. Of these, the question of optimum currency areas (OCA), advanced by Mundell (1961) and McKinnon (1963), is dominant in international macroeconomic policy discussions. However, consensus has not been reached about what are the essential ingredients of a common currency or a single currency area. For example, Meade (1956) implied that the necessary factor mobility should precede the creation of common currency, while Scitovsky (1958) argued that the creation of common currency would itself stimulate factor mobility. Frankel and Rose (1998) and Issing (2005) also indicated that causality could be either ways—should countries join a monetary union, the convergence indicators would be reinforced. In spite of the apparent contradiction between these two views, as Mundell (1961) argued, the concept of optimum currency areas is an empirical rather than a theoretical question. In the present context of increasing integration of China with East Asian production networks, that characterize vertical fragmentation of production processes across borders in East Asia, an attempt has been made in this paper to examine the costs to China's exports for not having a common currency in East Asia.

It is rational to argue that the development of East Asian production networks and increasing integration of China with these networks have created an environment with the much needed conditions to set the goal of monetary union in East Asia. These conditions include: (a) East Asian economies have synchronized business cycles, particularly after the Asian financial crisis, with key macroeconomic variables showing symmetric behavior throughout the business cycle in East Asia, (b) there is a substantial degree of factor mobility within the region and (c) there is a greater degree of interdependency among the countries in East Asia[1]. The literature generally suggests that capital has become highly mobile in East Asia since the early 1990s, while the condition of greater labor mobility is largely supplemented by vertical fragmentation of production processes across borders in East Asia. Unlike Europe, East Asian countries thus seem to have integrated their production and trading relationships in much a shorter period[2]. Fujita (2007; p. 6) remarked, "…it is surprising that, unlike EU and NAFTA, East Asia has attained such a high level of economic integration mainly through market mechanisms, with little support from region-wide political institutions." In fact, East Asian countries have not yet taken effective steps to move towards a monetary union, even though the region has attained economic integration to the level of Euro zone. The interesting question is as to what would happen to East Asian exports had there been a common currency area in East Asia. Particularly, this paper's focus is on China's exports which have raised concern in the U.S. and Europe. The paper assumes a hypothetical scenario of a common currency area in East Asia against which the impact of the existing

[1] For further details on these issues, see Kwan, 2001; Yusuf et al, 2004; Athukorala, 2005, McKinnon and Schnabl, 2003; Thorbecke and Yoshitomi, 2006; and Plummer and Wignaraja, 2007. The author thanks Ronald McKinnon for referring to synchronized business cycles in East Asia.

[2] East Asia has not only grown the fastest in the world (20.7 times over the 1970-2000 period), but also experienced deepening of economic interdependency within the region (the intra-regional trade share of East Asia has been approaching that of EU-15) (Fujita, 2007).

heterogeneous exchange rate policies of East Asian countries is evaluated. Though the present study focuses on China's exports to the world, the analytical framework is intended to produce useful implications for the creation of a monetary union in East Asia.

In the empirical trade literature, the conventional approach has been to estimate a gravity model by using a cross-country dataset, where the effect of currency union on trade, income and other macroeconomic variables is captured by a dummy variable (see, for example, Rose and van Wincoop, 2001; Frankel and Rose, 2002; Rose and Engel, 2002; and others). This approach is not suitable in the present study since the objective of the study is to quantify the cost of not having a common currency in East Asia. Another reason is that we study the problem, focusing only on China which largely features the terminal stage of the cross-border production networks in East Asia. An alternative conceptual framework is therefore developed in this study. The alternative conceptual framework is to include the interaction of the dynamic feature of the vertical integration of China's manufacturing with the production networks and the non-existence of a common currency area in East Asia directly into the model. The latter feature of the variable denotes the exchange rate policy asymmetry between China and other East Asian countries that supply intermediate goods to China. This is a time series variable. Two model specifications, one including the above variable along with other control variables and the other without the above variable but only with the control variables, are estimated to quantify the cost of not having a common currency area on China's exports.

The reminder of the paper is organized as follows. Section 2 describes background of China's surging exports to the rest of the world, particularly the U.S. and Europe, and its increasing linkage with East Asian production networks. Section 3 presents the conceptual framework setting out a model for the empirical estimations. Section 4 details on data and econometric methodologies. Section 4.1 introduces the dynamic panel data model, briefly highlighting the potential estimation problems. Section 4.2 discusses time-series properties of the observed data. Section 4.3 draws on estimation methods and specification tests. Section 4.4 describes data sources and main variables. Section 5 contains results and interpretation. Section 6 discusses robustness of the results. A final section brings the overall conclusions of this paper.

2. Production Networks in East Asia and China's Exports

China's export growth has been well documented in the literature[3]. The surge in China's exports has changed drastically the pattern of East Asia's trade surplus. Over the 1995-2005 period, the combined trade account balance of four major East Asian countries (i.e., Japan, South Korea, Taiwan, and China) grew from \$144 billions to \$491.7 billions as against the U.S. and EU-15 (see Table 1). But in this growing trade surplus, China's share accelerated from 34 percent to 66 percent. To the contrary, Japan's share declined from 60 percent to 23.5 percent. South Korea and Taiwan (hereafter called NIEs-2) also observed a declining trend, though much less than Japan, in their bilateral trading

[3] See, for example, Feenstra et al. 1998; Kawan, 2002; Eichengreen et al., 2004; Ahearne et al., 2006, Marquez and Schindler, 2006; and Rahman and Thorbecke, 2007.

relationships with the west. This indicates that China's spurring export growth has redefined the structure of the rest of East Asia's trade with the U.S. and EU-15.

However, China's phenomenal export growth happened in the context of its increasing integration with East Asian production networks. It is characterized by fragmentation of product value chain across borders in East Asia[4]. Here Japan and NIEs-2 organize those production processes that utilize on relatively higher skilled workers and produce sophisticated product prototypes, high-tech intermediate goods and capital equipments (Yusuf, 2003). These intermediate goods are transformed into finished products at assembly plants in China. The finished products are then exported throughout the world. Trade along these production networks, which is called vertical intra-industry trade (VIIT), has increased substantially over time (see Wakasugi, 2006; p. 25). China's integration with the East Asia production networks thus implies that the dollar cost of intermediate goods imported into China from the rest of East Asia represents a significant share of the '*gross value'* of Chinese exports to the U.S. and elsewhere. This point was emphasized in Greenspan (2005) and Lau and Stiglitz (2005).

Table-2 reflects on the vertical fragmentation of product value chain along the cross-border production networks in East Asia, by focusing on China's foreign trade in terms of product types and the country's major trading partners. The table summarizes China's trade statistics that are disseminated by Statistics Department of Customs General Administration of the People's Republic of China[5]. The export statistics are compiled in three categories: (a) the *"ordinary exports"* by local firms; (b) the *"processing exports"* by the foreign-owned firms (labeled *FDI-processing*) and (c) the *"other processing"* by Chinese owned firms. For the ordinary exports, local value addition constitutes the substantial portion of the *'gross value'* of those exports, whereas for the processing exports (both [b] and [c]), a larger share of the *'gross value'* originates in the upstream production blocks that are mainly located in Japan, NIEs and ASEAN. Feenstra and Spencer (2005, p. 1) remarked that processing exports were produced under contractual arrangements with foreign multinationals, where the ordinary exports did not have these arrangements. The role of VIIT is primarily related to the fragmented value chain of these processing exports.

Panel-A of Table-2 shows that in 2005, the *'processing imports'* by foreign-owned firms alone accounted for 42 percent of China's overall imports. Of this 42 percent, about 70 percent came from other East Asian countries. Similarly 50 percent of the *'other processing imports,'* which accounted for 16 percent of total imports, were also sourced from the rest of East Asia. Panel-B of Table-2 shows that in 2005, China's *'processing exports'* accounted for 55 percent of its overall exports. Of this 55 percent, about 76

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4 IBM's CEO, Sam Palmisano, noted in a recent *Foreign Affairs* article that an estimated 60,000 manufacturing plants were built by foreign firms in China alone between 2000 and 2003. Palmisano called it 'globally integrated business strategy' of multinational corporations. Fujita (2007, p. 18) argued that the spatial fragmentation of supply chain starting at the conception of the product and ending at its delivery had been the strategy of multinational firms (MNFs) to take advantage of difference in technologies, factor endowments or factor prices, and market sizes across countries.

5 These data are purchased by the Research Institute of Economy, Trade and Industry (RIETI) from China's Customs Statistics Information Center, Economic Information Agency, Hong Kong.

percent went to the Americas and Europe, if Hong Kong is arguably treated as an entrepôt of trade and transshipment to the west[6]. Thus the production network linkage of China with East Asia has led to its rapidly growing trade deficit against the rest of East Asia, but mounting bilateral trade surplus against the Americas and Europe, particularly the U.S. and EU-15.

Panel-C of Table-2 shows China's trade account balance in 1993 and 2005. In 2005, China incurred a deficit of $140 billions against Japan, NIEs-2 and ASEAN-5, but a surplus of about $290 billions against the U.S. and EU-15. The *'processing trade'* alone accounted for about 60 percent of China's bilateral trade deficit against East Asia, while another 35 percent occurred on account of '*other processing trade*,' which were also mostly related to the imports of capital equipments and industrial supplies intended for further processing by Chinese-owned firms. On the other hand, the *'processing trade'* with the U.S. and EU-15 alone accounted for about 83 percent of China's combined trade account surplus against those two end-user markets. The rest is being explained by China's trade in '*ordinary goods*.' This highlights the role of processing trade in the overall trade account balance of China. It further reflects on the degree of China's integration backward with the East Asian production networks and forward to the world trading system.

The present study, therefore, distinguishes between *'processing exports'* and *'ordinary exports'* for analyzing the impact of a common currency area on China's exports. It is believed that the existing heterogeneous exchange rate policies in East Asia will largely affect China's *'processing exports,'* by affecting the growing pattern of VIIT, which underlies the vertical organization of East Asian production networks. In order to do it, the study first aims to provide consistent estimates on the impact of appreciation of both the RMB exchange rate and the exchanges rates of all other East Asian countries that supply intermediate goods to China. The study then estimates potential costs to China's exports for not having a common currency area in East Asia. Accordingly, a conceptual framework has been developed in the following section.

3. Conceptual Framework of the Study

It is rational to argue that that 'gross value' of China's exports arises from the interlinked production networks in East Asia. In particular, the value chain of China's processing exports is vertically fragmented along the networks, giving rise to the vertical intra-industry trade (VIIT) amongst East Asian countries. Figure-1 depicts a schematic view of the fragmented value chain and its relationship with real exchange rate movements in East Asian countries. Japan, NIEs-2, ASEAN-4 and China are shown in four rectangular-shaped cells. x_j represents incremental real value added by production blocks in country j to the value chain of a product. We assume that the production value chain is ended in China. The solid line indicates the general pattern of intra-regional VIIT trade in East

[6] Hong Kong has yet remained as an entrepôt to facilitate transshipment of China's final exports to the rest of world, largely to circumvent both trade-related and non-trade barriers in the U.S. and EU-15 (Kwan, 2002; and Fung and Lau, 2001).

Asia. Since bilateral exports are recorded *'gross'* at every point of cross-border transfer, Σx_j's rather than the x_j's are observed. Therefore the real value of processing exports from China to country i is *'the gross'* Σx_j instead of x_C. The dashed line indicates the price-adjusted real exchange rates within East Asian countries and also between East Asian countries and country i. Note the locus of the RMB exchange rate in the network. Though the vertical integration of production processes along the networks has blurred independence of borders in East Asia, their exchange rate and monetary policies are rather independent and asymmetric from one another. The issue is about how a depreciation of country i's nominal exchange rate, say the quasi-global currency like the U.S. dollar or its new rival the Euro, will enter into the production networks and affect VIIT as well as final exports from China. The related hypothetical question is about what would happen, had there been a currency area arrangement in East Asia.

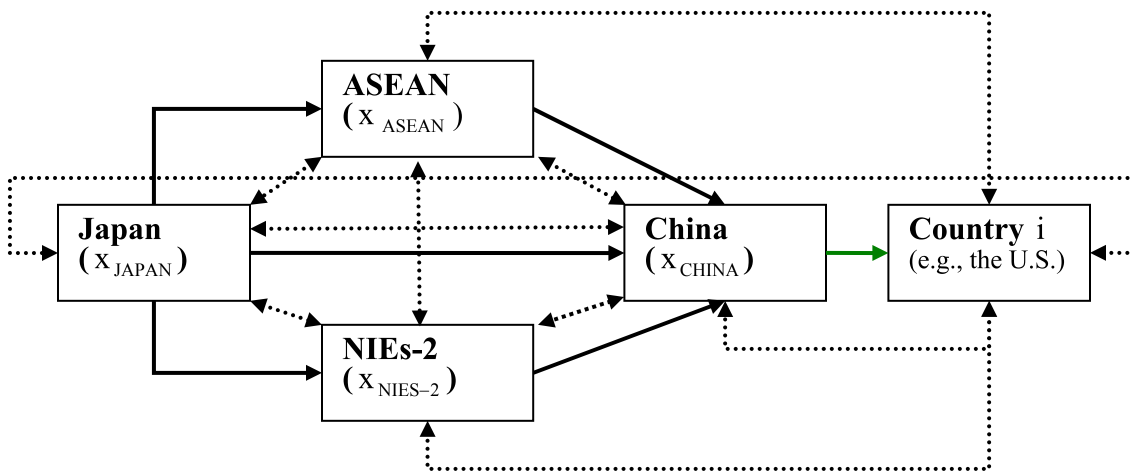

Figure 1: Schematic View of Production Networks and Real Exchange Rate Relationships

There are two possible scenarios in the actual policy environment. First the nominal exchange rate of country i may experience a discrete depreciation only against the Chinese RMB. Second, it may depreciate against all East Asian exchange rates including the RMB. In either of the above two cases, it should be recognized that nominal exchange rates and national monetary policies are mutually determined in financially open economies (see, Lahiri and Vegh, 2001; and Calvo and Reinhart, 2002). But East Asian countries do significantly differ in the application of these policy instruments (Ogawa and Ito, 2002; and Ogawa and Yang, 2006). Therefore to what extent a nominal appreciation of either the Chinese RMB or all the East Asian currencies, particularly against the world invoice currency, the U.S. dollar, will translate into a real appreciation of the respective currencies is unknown. However, if policy reaction by individual East Asian countries tends to be heterogeneous and uncoordinated, there will be significant flexibility in the real exchange rates within East Asia and between East Asia and the rest of the world. Its impact on the production networks and related VIIT will be highly reflected in the case of processing exports from China, but not so in the case of ordinary exports.

Three sets of real exchange rate relationships, as indicated by the dashed lines in Figure 1, are defined for the analytical purpose: (a) the real exchange rate between country j that supplies intermediate goods to China and country i that imports final exports from China, RER_{ji}; (2) the real exchange rate between China and country i, RER_{ci}, and (3) the real exchange rate between country j and China, RER_{jc} [7]. It is argued that difference in the reactions of the East Asian exchange rates to the depreciation of the rest of the world currencies, will create substantial variability in their real exchange rates. Price stability along the production networks will be distorted, thereby affecting growing supply-chain linkage of China's exports. Ogawa and Yang (2006; p. 17) find that since East Asian countries do not have any effective coordination mechanism in their international macroeconomic policies, they tend to resort to competitive devaluation. The policy of competitive devaluation is directly counterproductive to the operation of the production networks. How do we capture the effect of this exchange rate heterogeneity within East Asia?

As $RER_{ji} \equiv RER_{jc} * RER_{ci}$, one can not directly and separately measure the impacts of both the RER_{ji} and RER_{ci} variables on China's exports. This is because any empirical specification that includes both RER_{ji} and RER_{ci} as regressors is econometrically implausible. The present study instead recognizes that RER_{ji} can be factored into RER_{jc} and RER_{ci}. Moreover, the study conjectures that for China's processing exports whose value chain is fragmented vertically across borders in East Asia, while East Asian countries conduct asymmetric exchange rate policies, the flexibility in RER_{ji} will be mostly reflected in the misalignment of RER_{jc} and indirectly through its impact on RER_{ci}. Note that even if the RMB nominal exchange rate is fixed, changes in other East Asian exchange rates against the dollar, RER_{ji}, has the potential to misalign both the RER_{jc} and RER_{ci} variables which are the real exchange rate of East Asian country j against China and the RMB real exchange rate against country i respectively. The principal cause of this misalignment is their heterogeneous monetary and exchange rate policies, but its impact will be compounded by the degree of integration of China with the rest of East Asia. Let ω_j be the weight of country j in the *'gross value'* of Chinese final exports. Then the interaction term $\omega_j * RER_{jc}$ captures both the dynamic integration of China with country j and the asymmetry in actual policy making between China and country j. For the present analysis, a new variable is thus created as follows: $RER_w = \Sigma_j \omega_j RER_{jc}$. The term ω_j is proxied by the share of country j in China's

[7] For the home country i and foreign country j with price levels p_i and p_j and e_{ij} being the nominal exchange rate (in terms of home currency), we say that home country experiences a *real appreciation*, and the foreign country a *real depreciation,* when $RER_{ij} = p_i / e_{ij} p_j$ rises.

imports for processing. Here the subscript j represents the following East Asian countries: Japan, South Korea, Taiwan, Hong Kong, Indonesia, Malaysia, the Philippines, Singapore, and Thailand. These countries together supplied about 70 percent of China's processing imports in 2005. Note the variable RER_w is thus a time series variable in the panel of China's bilateral exports. The variable is defined as the weighted real exchange rate between East Asian countries that organize larger part of the fragmented (cross-border) value chain and China where final-stage assembly is done to produce "processing exports." Its flexibility is the unintended misalignment in relative prices between China and the rest of East Asia. A movement towards establishing a currency area in the region will first minimize and then eliminate this flexibility.

Having defined the variable of real exchange rate flexibility between China and the rest of East Asia, the study then considers two cases. One is that East Asian countries continue with their heterogeneous exchange rate and monetary polices, as their exchange rates appreciate against the rest of the world currencies, particularly the U.S. dollar. The other is a hypothetical case that there exists perfect coordination in East Asian exchange rate management, such as a common currency case. These two cases will jointly provide a framework to measure the potential costs to China's exports for not having a common currency in East Asia. Both of these cases are illustrated below.

Let China's export demand equation which approximates the true demand function be[8]

$$y_i = \beta_1 RER_w + \beta_2 RER_{ci} + u_i. \qquad (1)$$

Here β_1 measures the impact of relative price changes between China and other East Asian countries that supply intermediate goods to China, and β_2 measures the impact of relative price changes between China and country i that imports China's final exports. The variable RER_w represents interactions between the asymmetry in exchange rate policies between China and the rest of East Asia and the dynamic integration of China with the rest of East Asia, which underlies rapidly growing vertical intra-industry trade (VIIT) in the region. The VIIT essentially requires a symmetric policy making across countries that jointly organize such integrated networks. Therefore, the coefficient of the RER_w variable is the point estimate of the lack of correspondence of actual policy making from the desired symmetric case of having a common currency area. Let the hypothetical model, which ignores the influence of VIIT and the existence of asymmetry in exchange rate management in East Asia be

$$y_i = \beta_2^* RER_{ci} + v_i. \qquad (2)$$

In this case, the RER_w variable is excluded. Omission of RER_w variable will cause upward bias in the coefficient of RER_{ci}, because with the existing VIIT and asymmetric policies across East Asia, model (1) is the 'true' model from econometric estimation

[8] The set of other controls is excluded for clarity of the discussion.

point of view. The bias is $p\lim(\hat{\beta}_2^* - \beta_2) = \beta_1 b_{12} > 0$. Here $\hat{\beta}_2^*$ is the estimated upward-biased coefficient of RER_w, and b_{12} is the regression coefficient in the "auxiliary" regression of the excluded variable RER_w on the included variable RER_{ci} (see Maddala, 1977; p.156). A greater misalignment of real exchange rates of East Asian countries that organize upstream production processes of China's processing exports will tend to inflate β_1 and hence $\beta_1 b_{12}$. Now consider that there exists perfect coordination. In other words, there is dollar parity in East Asian countries. Then it implies that RER_{ci} is the only relevant variable and RER_w is the irrelevant variable in explaining China's exports behavior. Hence the hypothetical model

$$y_i = \beta_2 RER_{ci} + v_i \quad (2a)$$

would now become the true model. The least square estimator $\hat{\beta}_2 \to \beta_2$. In other words, $p\lim(\hat{\beta}_2 - \beta_2) = \beta_1 b_{12} \to 0$. This is because the irrelevance of RER_{jc} would cause $\beta_1 \to 0$.

Now the impacts of asymmetric exchange rate policies of East Asian countries and the exact measure of cost of pursuing such asymmetric policies can be estimated for China's exports. Since East Asian countries jointly organize a vertically fragmented production networks and hence conduct increasing volume of vertical intra-industry trade (VIIT), the impact of exchange rate flexibility on China's exports will be measured by the coefficients β_1 and β_2 in model (1). The latter term is the true parameter, rather than the biased RMB coefficient $\hat{\beta}_2^*$ of the misspecified model (2). Note that the RMB coefficient in the correct model is $\beta_2 = \hat{\beta}_2^* - \beta_1 b_{12}$, where $\hat{\beta}_2^* < 0$ and $\beta_1 b_{12} > 0$. Adverse effect of heterogeneous policy reactions on the RMB coefficient is therefore $\beta_1 b_{12}$. The cost of not having a common currency is measured by $\beta_1 + \beta_1 b_{12}$. In other words, if there is a common currency in East Asia and hence stable relative price relationships within East Asian countries, the coefficient $\beta_1 \to 0$ and thus "the enhanced effect" $\beta_1 b_{12} \to 0$. This means $(\beta_1 + \beta_1 b_{12}) \to 0$. Therefore the impact of asymmetric exchange rate policies in East Asian countries on the final exports from China is indicated by $\beta_1 + \beta_1 b_{12}$, which could be avoided had there been a common currency in East Asia.

This now clarifies how heterogeneous policy reactions and greater flexibility in East Asian real exchange rates will affect East Asian production networks and therefore China's exports. The above framework is applied in the empirical estimation for both the panels of 'processing exports' and the 'ordinary exports' from China.

4. Econometric Methodologies

4.1 The Dynamic Panel Data Model

In this study, China's export demand function is modeled in a dynamic framework which takes the general form of an autoregressive and distributed lag (ADL) model of order (2, 2)[9]:

$$y_{it} = \sum_{k=1}^{p} \alpha_k y_{it-k} + \boldsymbol{\beta}'(\mathbf{L})\mathbf{x}_{it} + \boldsymbol{\gamma}'\mathbf{z}_i + \eta_i + \boldsymbol{\delta}_i'\mathbf{d}_{it} + u_{it},$$
$$t = p+1, \cdots, T; \qquad i = 1, \cdots, N. \tag{3}$$

Here y_{it} represents China's bilateral real exports (either processing or ordinary) to country i, the vector $\mathbf{x}_{it} = [RER_{cit}\ RER_{wt}\ GDP_{it}]'$, is the set of right-hand side variables that can be either endogenous, predetermined, and/or truly exogenous, $\boldsymbol{\beta}'(\mathbf{L})$ is the coefficient vector of polynomials in the lag operator, RER_{cit} represents the bilateral real exchange rate between China and country i which imports final exports from China (an increase denotes an appreciation of the Chinese RMB), RER_{wt} represents the weighted real exchange rate between China and the countries that supply intermediate goods to China (an increase denotes an appreciation of real exchange rates of supplying countries against China), and GDP_{it} represents the real income of the importing country i. The vector $\mathbf{z}_i$ is a set of gravity variables such as, the distance between China and country i and dummy variables indicating whether the two countries are contiguous, share a common language, and have a colonial link[10]. The variables $(y_{it}, \mathbf{x}_{it})$ are measured in natural logs and vary both over time and across countries; while $\mathbf{z}_i$ only vary across countries. The model also includes fixed effect η_i, capturing unobserved factors that are not explicitly included as explanatory variables but affect the cross-sectional units of the sample and the values of the dependent variable observed for them. The vector d_{it} indicates the deterministic variables (intercept and/or trend terms) and δ_i indicates the corresponding vector of coefficients. The error terms u_{it} are assumed to be serially uncorrelated and distributed independently across cross-sectional units.

The dynamic specification (3) is intended to approximate China's true export demand function in the framework of imperfect substitutes model. However since the observed demand function is the equilibrating behavior of both the supply-side as well as the demand-side of the model, the relationship between quantities and prices is, at least in theory, simultaneous. The empirical trade literature has taken the supply-side by assumption that the price elasticity of supply is infinite (e.g., Maddala, 1977; and

[9] The dynamic feature of the model is related to the assumption that there are types of adjustment costs, such as transactions costs and/or that agents react only slowly to changes in their environment due to habit or inertia. Finite distributed lags are assumed to capture unobservable expectations about future outcomes (see, Hendry, Pagan and Sargan, 1984; pp. 1037-1039). The model can also be considered as a serial correlation model of Anderson and Hsiao (1982). However, we do not need to impose the implied common factor restrictions, and alternatively, the dynamics may be thought of as an empirical approximation to some more general adjustment process, as suggested by Blundell and Bond (1998).

[10] See Anderson, 1979; Deardorff, 1995; and Kalirajan, 1999 & 2007 for theoretical foundation of gravity equation and on the relevance of gravity variables in estimating bilateral trade equations.

Goldstein and Khan, 1985). Following Working (1927, p. 224), the present study however includes a set of gravity and/or deterministic variables as controls in order to correct for the influence of determinants which cause demand curves to shift. This is believed to produce a better approximation of the true demand curve, even though the original demand schedules fluctuated widely.

Conventional practice has been to assume that $(y_{it}, \mathbf{x_{it}})$ are unit root processes and that there exist cointegration relations. The assumption conveniently provides researchers a framework for modeling both the long-run equilibrium and the short-run dynamics. However Hylleberg and Mizon (1989, p.116) argued that an important criterion for econometric model adequacy is congruence of the model with time series properties of the observed data, embracing stochastic and/or deterministic trends for the non-stationary components, and appropriate representation of the temporal dependence of the stationary components. They argued, instead of assuming that there were cointegration relations, applied econometricians should obtain exact time series properties of the data. This is more important in the case of panel data, since their non-stationary characteristics are difficult to assess. Even the presence of non-stationarity in the data does not mean that the cross-sectional units are cointegrated and that the conditional distribution of the regression model would be stationary. We therefore assess time series properties of the data before embarking on any econometric estimation.

4.2 Time Series Properties of the Data

In order to know whether non-stationarity in the data is due to a deterministic time trend or unit root, panel unit root tests are conducted for each of the four principal variables of interest, i.e., y_{it}, RER_{cit}, RER_{wt}, and GDP_{it}. In theory, test of the null hypothesis of a unit root depends on whether or not deterministic elements (a constant term or a time trend or both) are included in the estimated regression and on whether or not the random walk that describes a true process includes a drift term. Since there is no specific null hypothesis about the process generating the individual data series, we take a general specification that can plausibly describe the data under both the null hypothesis and the alternative hypothesis. Following Levin, Lin, and Chu (2002), a univariate dynamic panel data model is formulated as follows[11]:

$$\Delta y_{it} = \alpha_i' d_{it} + \beta_i y_{it-1} + \sum_{L=1}^{P_i} \theta_{iL} \Delta y_{it-L} + e_{it} \qquad (4)$$

Here d_{it} indicates the vector of deterministic variables and α_i indicates the corresponding vector of coefficients. The null hypothesis is that each cross-sectional time series contains a unit root against the alternative hypothesis that each time series is stationary. In other words, the null hypothesis $H_0 : \beta_i = \beta = 0$, for all i, against the *homogenous alternative* $H_1 : \beta_1 = \cdots = \beta_i = \beta < 0$, for $i = 1,2,\cdots,N$. Levin et al. suggested

[11] The form is the Sims, Stock and Watson (1990) canonical form for higher order autoregressive processes, originally proposed by Fuller (1976).

a three-step procedure to implement the panel unit root tests. First, two auxiliary regressions are carried out to generate orthogonalized residuals. Second, the ratio of long run to short run innovation standard deviations $(s_i = \sigma_{yi} / \sigma_{\varepsilon i})$ for each cross-sectional unit is estimated. The estimate of the average standard deviation ratio $(\hat{S}_N = (1/N)\sum_{i=1}^{N} \hat{s}_i)$ is then used to adjust the mean of the $t-$statistic in the final step. In this final step, all cross-sectional and time series observations are pooled to estimate: $\tilde{e}_{it} = \delta\tilde{v}_{it-1} + \tilde{\varepsilon}_{it}$, where $\tilde{e}_{it}$ and $\tilde{v}_{it-1}$ are the normalized residuals estimated in step 1. The conventional t_δ statistic from the above estimation is then adjusted to derive adjusted t_δ^* statistics, which follow standard normal distributions. Breitung and Pesaran (2005) argued that Levin-Lin-Chu panel unit root tests have smallest size distortions and also perform best against the homogenous alternative and also when time dimension T is small.

Table A1 provides the results of the Levin-Lin-Chu panel unit root tests. The results indicate that real exports and real GDP are trend stationery series with first-order autoregressive error processes while the RMB real exchange rate is an I(0) stationary process with higher order autoregressive terms. By contrast, for the weighted exchange rate variable, RER_{wt}, being a time series variable, we obtain both the augmented Dickey-Fuller and the Philips-Perron unit root test statistics. The test statistics indicate that RER_{wt} is a unit root process, regardless the number of higher-order autoregressive terms and/or a drift term included in the estimated regression. Thus the general finding is that the dependent variable (real exports) is trend stationary, and the set of regressors includes one trend stationary series (real GDP), one stationary series (the RMB real exchange rate), all with serially correlated error terms. The variable RER_{wt} is a stochastic trending series with higher-order autoregressive terms. There is no one series that contains both deterministic and stochastic trend components. These results suggest that we cannot model the variables $(y_{it}, \mathbf{x}_{it})$ as a cointegrated system. The study also considers Pesaran (2003) cross-sectionally augmented Dickey-Fuller (CADF) regression that allows for panel heterogeneity. The null hypothesis is $H_0 : b_i = 0$, for all i, against the alternative $H_1 : b_i < 0$, for $i = 1,2,\cdots, N$. The results show that y_{it} and GDP_{it} are trend stationary series with serially correlated errors, while RER and WRER series are unit root processes. Again since y_{it} and GDP_{it} are evidently trend stationary series, it is unlikely that conditional distribution of any regression model with y_{it} as the dependent variable would be stationary. In other words, cointegration and error correction representation is not the plausible specification for the observed data.

4.3 Estimation Methods and Specification Tests

The empirical specification (3) introduced in section 4.1 is a dynamic error-component model. The composite error term is shown to be $v_{it} = \eta_i + u_{it}$. As Hausman (1978) argued, the unobserved fixed effects η_i in dynamic panel model are highly likely to be correlated with the observed exogenous variables and hence the model would be 'the fixed effects model,' rather than the uncorrelated random effects model. We do not conduct the Hausman test in choosing between random effects and fixed effects approach. It is neither necessary because the Hasuman test assumes strict exogeneity of $\mathbf{x}_{it}$ variables, while the present study essentially allows the vector $\mathbf{x}_{it}$ to include a set of exogenous, predetermined and/or endogenous variables. However it is important to recognize that in dynamic panel model that includes unobserved fixed effects, the pooled OLS estimators are upward biased They are based on the restrictive assumptions that $E(\mathbf{x}'_{it}u_{it}) = 0$ and $E(\mathbf{x}'_{it}\eta_i) = 0$, for $t = 1, \cdots, T$. The dynamic model with lagged dependent variable in $\mathbf{x}_{it}$ must violate the assumptions because y_{it-1} and η_i are correlated.

Nickell (1981) showed that standard methods of eliminating the fixed effects, such as first-difference or within estimation, would also lead to seriously biased measures of coefficients and therefore covariance estimators are inconsistent. For an autoregressive model that includes a vector of truly exogenous variables, Nickell showed that the within estimation of the autoregressive parameter would be downward biased. Whereas, the bias in the coefficients of the included exogenous variables would depend on the relationship between the exogenous variables and the lagged dependent variable y_{it-1}. If an exogenous variable is positively correlated to y_{it-1}, then the coefficient estimate would be upward biased and vice-versa (Nickell, 1981; p. 1424). Wooldridge (2002) shows that if u_{it} are correlated with future values of the explanatory variables in the sense that $E(\mathbf{x}'_{it}u_{is}) \neq 0$ for $s < t$, the strict exogeneity assumption fails in a dynamic panel model. And this will cause unknown bias in the fixed effect estimator. In addition, if the process $\{\mathbf{x}_{it}\}$ has very persistent elements, the within estimator can also have substantial bias.

With the inconsistency of the covariance estimators in dynamic panel data model, Anderson and Hsiao (1982) showed that the maximum likelihood estimators were also highly sensitive to alternative assumptions about initial conditions and asymptotic plans, i.e., in the way the time series dimension T or the cross-sectional dimension N tends to infinity. As a result they proposed to estimate the model in first-differences by instrumental variables using either y_{it-2} or Δy_{it-2} as instruments. Anderson and Hsiao argued that the advantage of these estimators was that they were consistent whatever the form of the initial conditions and whether T or N or both were tending to infinity. Arellano and Bond (1991) showed that in the dynamic model which includes exogenous variable, there are values of α and $\boldsymbol{\beta}$ between 0 and 1 for which the instrument matrix $Z_i^* = \mathrm{diag}(y_{it-2})$ may not be valid, causing the estimators to be inefficient.

In the tradition of Holtz-Eakin, Newey and Rosen (1988), Arellano and Bond (1991) suggested the Generalized Method of Moments (GMM) approach to exploit further population moment conditions in the first-difference equations of the dynamic specification. The basic assumption of the model is that u_{it} have finite moments and in particular, $E(u_{it}) = E(u_{it}u_{is}) = 0 \quad \forall t \neq s$. That is, u_{it} are assumed to be serially uncorrelated. In the first-difference equations, the above assumptions lead to a set of linear orthogonality conditions. The complete set of moment conditions, for the present case of ADL(2,2) model, has the form of $E(Z_i' \Delta u_i) = 0$ for $i = 1, \cdots, N$, where $\Delta u_i = (\Delta u_{i4}, \cdots, \Delta u_{iT})'$. Here Z_i is the matrix of valid instruments in the first-differenced equations. However Z_i crucially depends on whether the vector x_{it} includes endogenous, predetermined or strictly exogenous variables. The present study does not impose any arbitrary restrictions as to exogeneity of the included variables, rather allows the model be flexible and take the properties of the observed data.

Arellano-Bond GMM first-difference estimators that are based on the moment conditions that $E(Z_i' \Delta u_i) = 0$ are weakly identified when the instruments are weak in the sense that they have a low correlation with the included endogenous variables (Maddala and Jeong, 1992; Blundell and Bond, 1998; and Stock and Wright, 2000). Blundell and Bond also showed that GMM first-difference estimators can be seriously downward biased in two important cases. First, as the value of α approaches to unity, and second, as the relative variance of the fixed effects η_i, i.e., $(\sigma_\eta^2 / \sigma_v^2)$ increases to infinity. Following Arellano and Bover (1995) that lagged differences can be used as possible instruments for equations in levels, Blundell and Bond considered another $(T - p - 1)$ linear moment conditions: $E(u_{it} \Delta y_{it-1}) = 0$ for $t = p + 2, \cdots, T$, where p denotes the autoregressive order of the model. Based on a stacked system comprising all $(T - p - 1)$ equations in first differences and the $(T - p - 1)$ equations in levels corresponding to $t = p + 2, \cdots, T$, Blundell and Bond suggested the following extended instrument matrix $Z_i^+ = \text{diag}(Z_i \vdots \Delta y_{i1} \cdots \Delta y_{is})$ for $t = p + 2, \cdots, T$ and $(s = 1, \cdots, T - 2)$. The estimator based on this extended instrument matrix is called GMM system estimator. For both the cases of GMM first-difference and GMM system estimators, the alternative choice of the weighting matrix $\mathbf{A_N}$ will produce one-step or two-step estimators (for further details on the weighting matrix, see Hansen, 1982).

Since the weak instrument problem is very likely in the macroeconomic time series data, we assess persistency characteristics of each individual time varying data series included in the model (3). Appendix-1 details on the estimation procedure for assessing persistency in $(y_{it}, \mathbf{x_{it}})$. The exercise is useful as this study allows for potential endogeneity among the variables. We obtain consistent estimates of persistency for $(y_{it}, \mathbf{x_{it}})$. The results show that the dependent variable y_{it}, though a trend stationary process, tend to be highly persistent. RER_{cit} and RER_{Wt} appear to be moderately persistent, while GDP_{it} posits strong persistency. Table A2 provides detailed results on persistency in the observed data. The findings on persistency characteristics in the observed data indicate that not only are the pooled OLS and the fixed effects estimators inconsistent, but so too are the GMM-diff estimators that are likely to be weakly identified. The study thus follows Arellano and Bover (1995) and Blundell and Bond

(1998) and derives GMM system estimator which is based on an extended instrument matrix.

This paper estimates two benchmark specifications of model (3) in order to estimate the costs to China's exports of asymmetric exchange rate policies in East Asia as against the hypothetical case of a common currency. The benchmark specifications are:

$$y_{it} = \sum_{k=1}^{2} \alpha_k y_{it\text{-}k} + \beta_0 GDP_{it} + \beta_1 GDP_{it-1} + \beta_2 GDP_{it-2} + \xi_0 RER_{cit} + \xi_1 RER_{cit-1} + \xi_2 RER_{cit-2} + \psi_0 RER_{wt} + \psi_1 RER_{wt-1} + \psi_2 RER_{wt-2} + \boldsymbol{\gamma}'\mathbf{z_i} + \eta_i + \boldsymbol{\delta}_\mathbf{i}'\mathbf{d_{it}} + u_{it} \quad (3.1)$$

$$y_{it} = \sum_{k=1}^{2} \alpha_k^* y_{it\text{-}k} + \beta_0^* GDP_{it} + \beta_1^* GDP_{it-1} + \beta_2^* GDP_{it-2} + \xi_0^* RER_{cit} + \xi_1^* RER_{cit-1} + \xi_2^* RER_{cit-2} + \boldsymbol{\gamma}^{*\prime}\mathbf{z_i} + \eta_i + \boldsymbol{\delta}_\mathbf{i}^{*\prime}\mathbf{d_{it}} + v_{it} \quad (3.2)$$

The specification (3.1) is the true model, while the specification (3.2) is the hypothetical model based on the counterfactual assumption of a common currency area in East Asia. Appendix-2 provides specific information regarding the orthogonality conditions and the related GMM extended instrument matrices that have been actually used to obtain the parameter estimates. In particular, we allow real exchange rate variables to be strictly exogenous, predetermined or endogenous and accordingly define the corresponding instrument matrices. We thus obtain three sets of GMM estimators for both the specifications. In addition to the GMM-system estimators, we also obtain pooled OLS and fixed-effect estimator in each case in order to show relative performance of the GMM estimators.

The recent empirical trade literature suggests that empirical export demand equation should include a supply shift variable. Mann and Plück (2005) argued that exporters' ability to produce more variety with increasing returns to scale would cause shifts in export supply curve for the exports of fast-growing countries such as China. Chinn (2005) argued that the empirical trade model should include a supply shift variable; otherwise the model would imply a constant relationship between domestic consumption of exportables and production of exportables. Following the recent empirical trade literature (e.g., Helkie and Hooper, 1988; Bayoumi, 1999; and Chinn, 2005), the present study thus augments the benchmark specifications by including alternative proxies to control for increased capacity of exporters to supply more variety. The purpose is to check robustness of the parameter estimates of the benchmark model.

Another potential debate is on the use of an appropriate deflator for the case of China's exports. Following the recent literature, we apply Hong Kong's export price index, the U.S. import price index of manufactured imports from non-industrial countries, and the U.S. CPI as the alternative deflators to obtain real measures of China's exports data that are reported in current U.S. dollars. The deflators are further explained in the data section. Again the motivation is to check robustness of the parameters of interest to the use of alternative deflators.

Since our sample has $N = 33$ and $T = 14$, we use less than the available valid moment restrictions in order to avoid the problem of overfitting the instrumented variables and thereby causing the results biased towards those of OLS[12]. Sargan (1958) and Amemiya (1977) suggest that from the standpoint of obtaining desirable small sample properties, one should try to conserve the number of orthogonality conditions used in the GMM estimation (Hansen, 1982; p. 1035). Following Roodman (2006), the present study also collapse the "GMM-style" moment conditions into groups and sums the conditions in each group to form a smaller set of moment conditions. Since standard errors of two-step GMM system estimator tend to be severely downward biased, we apply a finite sample correction to the two-step covariance matrix as suggested by Windmeijer (2005) and thereby obtain corrected standard errors estimates.

Finally, the study provides the standard specification tests. Let $\overline{u}_{it}$ be the first differences of serially uncorrelated errors (u_{it}). Then $E(\overline{u}_{it}\overline{u}_{i(t-1)})$ need not be zero, but the consistency of the GMM estimators fundamentally depends upon the assumption that $E(\overline{u}_{it}\overline{u}_{i(t-1)}) = 0$. We thus report both $m1$ and $m2$ tests for first-order and second-order serial correlation in the first-differenced residuals, asymptotically distributed as N(0,1) under the null of no serial correlation. They both are reported in order to discriminate the situation if the errors in levels follow a random-walk process from the situation if the errors in levels are not serially correlated. Next we provide Sargan/Hansen test of over-identifying restrictions. When the number of orthogonality conditions (r), exceeds the number of parameters to be estimated (k), estimation of the model parameters sets k linear combinations of the r sample orthogonality conditions equal to zero, at least asymptotically. Thus when the model is true, there are $(r-k)$ linearly independent combinations of the orthogonality conditions that ought to be close to zero but are not actually set to zero (Hansen, 1982). These linear combinations of sample orthgonality conditions are used to obtain Hansen J statistic. Hansen J statistic is thus a test of the over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. For both one-step robust estimation (and also for two-step estimation), the Hansen J statistic is the minimized value of the two-step GMM criterion function and is asymptotically valid test statistic of the model restrictions.

4.4 Data[13]

China's Disaggregated Trade Flows: The study uses annual data on China's bilateral exports and imports statistics, disaggregated into ordinary and processing categories, vis-à-vis a panel of 33 countries over the 1992-2005 period. The data are compiled by the Statistics Department of Customs General Administration of the People's Republic of China and published by the Economic Information and Agency, Hong Kong.

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[12] To check the parameter stability, we also exploit all the valid moment restrictions.

[13] The author is grateful to the Research Institute of Economy, Trade and Industry (RIETI) for providing the datasets and other research supports.

Deflators: Since China does not report export price index, we follow the recent empirical literature and apply three alternative deflators, i.e., Hong Kong export price index, the U.S. consumer price index, and the U.S. import price index of manufactured imports from non-industrial countries. Liang and Fung (2005) argue that Hong Kong price index is highly likely to trace the movement of export price of China's exports. The reasoning is the traditional role Hong Kong has played as an entrepôt to transship China's exports to the rest of the world. The data are taken from IMF International Financial Statistics, line 74d. Eichengreen, Rhee and Tong (2004) applied the U.S. consumer price index to deflate U.S. dollar imports. Thorbecke (2006) also applied the same deflator arguing that the measure would be appropriate if the bundle of goods and services exported from China corresponds to the bundle purchased by U.S. consumers. Cheung, Chinn and Fujii (2006) used the U.S. Bureau of Labor Statistics (BLS) price deflator for imports from non-industrial countries to deflate dollar value of China's exports. They find that this series closely matches the BLS price deflator for imports from China, which has been compiled since 2003.

Real Exchange Rates: Bilateral real exchange rates between China and country i, i.e., the RMB real exchange rate, RER_{ci}, and bilateral real exchange rates between China and country j that supplies intermediate goods to China RER_{jc}, are taken from the CHELEM database developed by Centre d'études prospectives et d'informations internationales (CEPII), the France's leading institute for research on the international economy. The weighted real exchange rate between China and countries that supply intermediate goods to China is defined as $RER_{wt} = \Sigma\omega_{jt}RER_{jct}$. Here ω_{jt} denotes the weight of country j in the 'gross value' of China's final exports and is proxied by the share of the country in China's imports for processing. The term ω_{jt} is annually updated over the sample period. Note that RER_{wt} is a time series variable and hence uniform across cross-sections in the panel of China's exports.

Real Output and Gravity Variables: Real income in the importing country i and a set of the gravity variables are also taken from CEPII. The gravity variables include distance and dummy variables indicating whether the two countries are contiguous, share a common language, and have a colonial link.

Proxies for the supply-side effect: To control for exporters' increased capacity to supply new varieties, this study uses several alternative proxies of the variable. They include real GDP of China (IFS series 99B_P), cumulative inward FDI to China (IFS series 78BED; the data is taken from McKinnon and Schnabl, 2007; p. 7), and China's fixed capital formation (IFS series 93_E).

5. Results and Interpretation

Tables 3 and 4 present estimates of the benchmark model for China's processing exports and ordinary exports respectively. The estimated equation is an empirical "gravity" equation to approximate true demand curve for China's exports. The purpose is to "correct" influences of determinants that cause the demand curve to shift and hence to derive consistent estimates of the parameters. The parameters of interest are coefficients of two real exchange rate (RER) variables. We do not report estimates of the coefficients of the gravity variables. Their point estimates are often statistically insignificant, but the variables are jointly significant. Real exports data in both cases are obtained by using Hong Kong export price index as the deflator[14]. In each table, the pooled ordinary least squares (OLS) and fixed-effect (FE) estimates are reported in the first two columns, which are followed by two GMM estimates[15]. GMM1 assumes that both the RER variables are exogenous in the model, while GMM2 assumes that they are predetermined[16]. Since the OLS estimates are upward biased and the FE estimates are downward biased, they provide a range within which the true autoregressive parameter should exist. A better approximation of the true autoregressive parameter would also correct potential bias in the estimates of other parameters of the model. This therefore provides the first criteria to judge relative consistency of our GMM estimates. All the GMM estimates that are reported here are one-step system GMM estimators, for which we believe inference based on the asymptotic variance matrix to be more reliable. Two-step system GMM estimators, though not reported here, provide very similar estimates of the parameters. All the tables report only parameters of interest and relevant specification tests. In general, the GMM estimates are relatively consistent and more efficient than the OLS and FE estimates. We find that both GMM1 and GMM2 specifications provide better approximation of the autoregressive parameter compared to the range implied by the OLS and FE estimates. The extended set of moment restrictions is not rejected by the Sargan/Hansen test of over-identifying restrictions. The tests of serial correlation in the first-differenced residuals are in both cases consistent with the maintained assumption of no serial correlation.

Table 3 shows estimation results of the equation for China's processing exports. Since these exports are manufactured along the inter-linked production networks in East Asia, their value chain is sharply fragmented across borders. According to the preferred GMM1 and GMM2 estimates, the elasticity of RER_w is in the range of −1.09 and −1.31. By contrast, the corresponding elasticity of RER_{ci} is roughly −0.75.The findings indicate that a 10 percent real appreciation of exchanges rates of all East Asian countries that supply intermediate goods to China will cause China's processing exports to decline by 11-13 percent. By contrast, a 10 percent unilateral RMB appreciation against the rest of the world currencies will cause China's processing exports to decline by about 7.5 percent.

[14] In the robustness analysis presented in the following section, we discuss further estimation results that are obtained by using alternative deflators as suggested by the contemporary literature and also by including alternative proxies to control exporters' increased capacity to supply new varieties.

[15] We report neither maximum likelihood estimates nor Anderson-Hsiao IV estimates for the reasons we explained in Section 4.3.

[16] The GMM estimates that are based on moment condition that arise from the assumption of endogenous RER variables are found to be poorly identified and further downward biased than the FE estimates. We therefore do not report them. Note that potential endogeneity of a variable x_{it} requires that its lags dated $(t-2)$ or earlier can only be the instruments in dynamic panel model that includes unobserved fixed effects. If x_{it} is, in fact, not endogenous, the assumption is restrictive and causes weak instrument problem.

Note that a real appreciation of exchange rates of the rest of Asia does misalign their relative price relationships with China. The RER_w coefficient essentially measures the impact of this misalignment. The flexibility in relative prices within East Asia arises from asymmetric policy reactions by individual governments in the region. It negatively affects China's processing exports because they are manufactured along East Asian production networks where China features the terminal stage.

Table 4 shows corresponding estimation results for China's ordinary exports. As we mentioned earlier, these ordinary exports are produced primarily by using local inputs. In other words, the role of East Asian supply chain is much less important in this case. Again we find GMM1 and GMM2 to be the preferred estimators. The results show that a 10 percent appreciation of RER_w will affect China's ordinary exports by about 6 percent, whereas an equivalent 10 percent of appreciation of the RMB will affect China's ordinary exports by about 8.5 percent. This indicates that a unilateral RMB appreciation would have larger impact on China's ordinary exports. It is because the extent of local value addition appears to be substantial in *'the gross value'* of China's ordinary exports. A relatively lower magnitude of RER_w coefficient signifies its weak link with East Asian supply chain.

What are the implications of the results regarding East Asian production networks and China's exports? Earlier we observed that *'processing trade'* is at the heart of China's integration with East Asia and also to the world trading system. The results in Table 3 indicate that a generalized appreciation of exchange rates of all East Asian countries that supply intermediate goods to China does affect China's processing exports by misaligning the relative price relationships between China and the rest of East Asia. This is what the coefficient of RER_w variable implies. This study further shows that real exchange rate misalignment within East Asia has a larger impact on China's processing exports than a unilateral RMB appreciation. This is because an RMB appreciation would affect only Chinese value added, whereas a relative price misalignment between China and the rest of East Asia would affect the dollar costs of intermediate goods imported into China from the rest of East Asia. The flow of these intermediate goods, the costs of which represent a larger share of the gross value of China's processing exports, is characterized by the vertical intra-industry trade (VIIT) along the production networks. The heterogeneous exchange rate policies in East Asia thus adversely affect the VIIT and growing pattern of cross-border fragmentation of production processes in East Asia. Since other East Asian countries too have integrated with these regional production networks, East Asia as an optimum currency area would essentially eliminate the adverse impact of asymmetric exchange rate and monetary policies on East Asian exports.

Recall that our main objective has been to quantify the costs of not having a common currency in East Asia for China's exports, particularly the processing exports that have strong production network-linkage with the rest of East Asia. The consistent estimates that are reported in Tables 3 and 4 are an intermediate step to that end. To focus on the issue of not having a common currency, we refer to Tables 5 and 6. These tables report estimates of the hypothetical model which is based on the counterfactual assumption that there exists a common currency, i.e., East Asian countries share a single currency. This causes the RER_w variable to be irrelevant as an explanatory variable.

In Table 5, we find that the coefficient of RER_{ci} is now upward biased by the magnitude of 0.20 from the consistent estimate of the correct model (i.e., −0.75). This upward bias arises from the omission of the relevant RER_w variable. We argue that if there is greater heterogeneity in exchange rate policy in East Asia, RER_w will have larger impact on China's processing exports by two mechanisms. One is its direct impact which is $\beta_1 = -1.09 \sim -1.30$, while the other is its indirect impact on the coefficient of RER_{ci} which is $\beta_1 b_{12} = -0.20$[17]. Had there been a common currency in East Asia, $\beta_1 \to 0$ and hence $\beta_1 b_{12} \to 0$. This indicates that the cost of not having a common currency in East Asia for China's processing exports is $\beta_1 + \beta_1 b_{12} = -1.29 \sim -1.50$[18]. The first component is the RER_w coefficient which measures the impact of real exchange rate misalignment within East Asia, while the second component is its indirect effect on the exchange rate of individual East Asian country. Here the country is China and the effect is negative to the extent of -0.20. However the cause is not the appreciation, rather the asymmetric reactions to the appreciation. Hence only if there is increasing coordination among East Asian countries to keep their relative price relationships stable and ultimately to form a monetary union, will the cost of existing asymmetric exchange rate policies first decline and then disappear.

The cost of not having a common currency area in East Asia is modest in the case of China's ordinary exports. Table 6 shows that the elasticity of RER_{ci} in the hypothetical model is upward biased by the magnitude of 0.10. Therefore the combined costs of not having a common currency area for the ordinary exports are $\beta_1 + \beta_1 b_{12} = -0.55 \sim -0.78$. This is just half of the costs we have estimated in the case of China's processing exports.

The results are indicative of the present state of East Asian monetary and exchange rate policies that are largely heterogeneous to external shocks. The deepening integration of East Asian countries with China has required similar integration and coordination in monetary and exchange rate policies in East Asia. But the present study shows that this has not been the case. As a consequence, a greater integration in production and trading relationships within East Asia appears not to be robust, rather very susceptible to external shocks such as a global economic slow down or a discrete/gradual appreciation of either the Chinese RMB or all East Asian exchange rates.

Though the present study is primarily intended to quantify the impact of heterogeneous exchange rate polices in East Asia on China's bilateral exports to the world, we also find

[17] Note that notations β_1, β_2 and b_{12} discussed in this section are consistent with the notations introduced in the conceptual framework of the paper, not the coefficients of the estimated model shown in the footnotes to the tables of estimation results.
[18] We estimate standard error of $\beta_1 + \beta_1 b_{12}$ by using delta method and find they are significant at 1 % level.

that external demand has been the single most important factor to stimulate China's exports. Income elasticity for China's processing exports is about 2.60. To the contrary, income elasticity of China's ordinary exports varies from 1.56 to 1.77. The findings reflect that China's processing exports are technologically sophisticated with higher income elasticity. On the other hand, the ordinary exports tend to be labor-intensive consumer goods for which income effects are naturally very low. A very high magnitude of income elasticity for the demand of the processing exports indicate that continued growth in global demand and China's integration with the world trading systems in 2001 appear to have accelerated China's exports to the rest of the world. To put it differently, a downturn outside of Asia has thus the potential to cause a large drop in China's processing exports. Since they are produced along the interlinked production networks in East Asia, such a global slow down would adversely affect intra-regional production and trading pattern, reducing employment and output throughout the region.

6. Robustness Analysis

Edward Leamer (1983) has argued persuasively that because any econometric analysis involves numerous debatable decisions, findings cannot be convincing unless they are shown to be robust. In the previous section, we already discussed robustness of our estimates to changes in several of the modeling decisions. Here we provide further robustness of parameter estimates to alternative variable definitions and the inclusion and/or exclusion of additional control variables in the estimates model.

The first major controversy centers on the use of an appropriate export price index to deflate China's nominal dollar exports to the real values. In the previous section, we consistently use the Hong Kong export price index as the deflator to define the dependent variable, China's real exports. Now we use two other alternative deflators, the U.S. CPI and the U.S. import price index of manufactured imports from non-industrial countries. For brevity, we report only the range of estimates of the parameters of interest[19]. For the case of China's processed exports, the coefficient of the RMB real exchange rate (RER_{ci}) is in the range of −0.71 to −0.79, while that of the weighted RER between China and all East Asian countries that supply intermediate goods to China (RER_{wt}) is in the range of −1.05 to −1.33. On the other hand, the corresponding results for China's ordinary exports show that the coefficient of RER_{ci} lies in the range of −0.83 to −0.90, while that of RER_{wt} lies in the range of −0.40 to −0.65. All these estimates are statistically significant at any reasonable level. Overall all these estimates are within one standard error of the corresponding estimates for the base case GMM estimates as they are reported in Tables 3 and 4 respectively. In other words, the estimates of the impact of heterogeneous of exchange rate policies and also of the costs of not having a common currency area on China's exports do remain robust, regardless of how we deflate nominal dollar value of China's exports. Next we augment the benchmark models by including a supply shift variable to control for structural break in the estimated relationships and to

[19] All the detailed results are available from the author. The range of estimates includes about 15 estimates depending on the variation in the definitions of the dependent variable and also in the specifications and estimations of the model.

examine if the estimates of the key parameters are robust to the inclusion of the variable. We discuss the results only for the panel of processing exports.

Since it is difficult to find a good proxy for the supply shift effect, we follow the recent trade literature and use two alternative proxies such as real GDP of China (Bayoumi, 1999), and China's gross fixed capital formation. The latter proxy is believed to be more reliable compared to the China's capital stock variable, which is likely to be measured with much error (see, for example, Feldstein and Horioka, 1980; and Chinn, 2005). The results show that if the model is augmented by including China's GDP, the GMM estimates of the coefficients of RER_{wt} and RER_{ci} variables in the panel of processing exports are −2.15 and −0.65 respectively. On the other hand, if the model is augmented by China's gross fixed capital formation, the corresponding estimates are −1.90 and −0.60 respectively. The implication of the augmented model is that our benchmark estimates in Tables 3 and 4 provide a conservative lower limit of the estimate of key parameters. In other words, the estimates of the costs of not having a common currency arrangement in East Asia are found to be conservative and robust. A more valid proxy of supply-side effect on China's exports would produce more consistent estimates of the parameters of interest. But it is evident that the actual costs of not having a common currency arrangement in East Asia, particularly for the processing exports, would be greater than the lower-bound benchmark estimates reported in the present study.

In the view of a small sample involving 34 cross-sections and 14 time periods, we consistently estimated the autoregressive and distributed lag model of second order. However, we found that in most cases the more distant lags of the dependent variable, real exchange rates, and real GDP did not cause any significant changes in the estimates of the key parameters, nor in their asymptotic efficiency. Following Grossman and Levinsohn (1989), we also tested separately for the joint significance of different lag lengths based on nested-hypothesis testing. The results provide general support for the ADL(2,2) model used in the present study.

7. Concluding Remarks

In this paper, we have developed a framework for assessing the impact of not having a common currency area in East Asia on East Asian production networks and China's exports behavior. The framework is built on two pillars. First a new variable is created in order to capture interactions between the increasing integration of China with the rest of East Asia and the asymmetry in exchange rate and monetary policies between China and other East Asian countries that supply intermediate goods to China. The other is the introduction of a hypothetical but estimable econometric model, which is based on a counterfactual assumption that there is a common currency in East Asia.

We apply this framework to the observed data on China's bilateral exports, divided into processing and ordinary categories, to a panel of 33 countries over the 1992-2005 period. The results show that the cost to China's processing exports for not having a common currency is more than double the corresponding cost to China's ordinary exports. The magnitudes of these costs are consistent with the hypothesis that a common currency in East Asia would further integrate East Asian production networks and promote those exports whose value chains are spatially fragmented across borders in East Asia.

However, the framework developed here warrants further refinement. It should be applied to a broader panel dataset of bilateral trade flows of all East Asian countries. Also, recent advances in exchange rate and trade literature might allow a better specification to derive more valid inference on the plausibility of a common currency in East Asia.

Table-1
Combined trade account balance of East Asian countries against the U.S. and EU-15: 1995-2006q2

	In billions of US dollars					In percent				
Country Year	Japan	China	South Korea	Taiwan	East Asia	Japan	China	South Korea	Taiwan	East Asia
1995	87	49	-3	12	144	60.1	33.8	-2.1	8.2	100
1996	69	59	-8	16	135	50.9	43.5	-6.0	11.5	100
1997	83	74	-3	16	169	49.2	43.5	-2.0	9.3	100
1998	103	85	15	22	224	45.8	37.7	6.8	9.7	100
1999	112	101	16	25	254	44.3	39.8	6.1	9.8	100
2000	120	125	20	27	293	41.1	42.7	6.9	9.2	100
2001	97	124	18	25	265	36.7	46.9	6.9	9.5	100
2002	94	148	18	23	283	33.4	52.3	6.3	8.0	100
2003	96	187	21	25	329	29.2	56.8	6.5	7.5	100
2004	109	250	33	25	417	26.2	60.1	7.8	5.9	100
2005	115	323	29	25	492	23.4	65.7	5.9	5.0	100
2006-Q2	54	144	12	11	221	24.5	65.1	5.3	5.2	100

Notes: East Asia includes Japan, South Korea, Taiwan and China. EU-15 represents Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and United Kingdom. Data for 2006-Q2 represents recent trade statistics of the U.S. (Jan. to June, 2006) and EU-15 (Jan. to March) in 2006. US dollar value of EU-15 Imports statistics prior to 1999 are estimated by applying the ECU exchange rate.

Sources: (a) U.S. Census Bureau, Foreign Trade Division [Website: http://www.census.gov/foreign-trade/balance/index.html#], and (b) EUROSTAT External Trade Database at HS-6 digit level [Website: http://epp.eurostat.ec.europa.eu/portal/].

Table-2

Panel-A: China's Imports – 1993 and 2005 (in %)

Partner Import categories	World	Japan (1)	S. Korea & Taiwan (2)	ASEAN-5 (3)	Hong Kong (4)	East Asia (5=1+2+3+4)	United States	EU-15	Rest of the World
1993									
Total imports	100.0	22.4	17.6	5.8	10.0	55.8	10.3	15.1	18.8
Ordinary imports	36.6	7.9	2.1	3.3	1.1	14.3	5.1	7.9	9.3
Processing imports	35.0	7.7	10.8	1.9	7.0	27.3	1.9	1.7	4.0
Other processing imports	28.4	6.8	4.7	0.7	2.0	14.1	3.3	5.5	5.5
2005									
Total imports	100.0	15.2	23.0	10.9	1.9	50.9	7.4	10.7	31.1
Ordinary imports	42.4	5.4	5.7	3.1	0.5	14.8	3.9	6.4	17.3
Processing imports	41.5	6.9	14.4	5.7	1.2	28.1	1.9	1.8	9.7
Other processing imports	16.1	2.9	2.9	2.1	0.1	8.1	1.5	2.4	4.1

Panel-B: China's Exports – 1993 and 2005 (in %)

Partner Export categories	World	Japan (1)	S. Korea & Taiwan (2)	ASEAN-5 (3)	East Asia (4=1+2+3)	Hong Kong	United States	EU-15	Rest of the World
1993									
Total exports	100.0	17.2	4.7	5.1	27.0	24.0	18.5	13.3	17.2
Ordinary exports	47.1	9.8	2.7	3.6	16.0	9.6	5.8	6.8	8.9
Processing exports	48.2	7.3	2.0	1.4	10.7	14.0	12.7	6.5	4.3
Other processing exports	4.7	0.1	0.0	0.1	0.2	0.4	0.0	0.0	4.0
2005									
Total exports	100.0	11.0	6.8	6.3	24.1	16.3	21.4	17.3	20.9
Ordinary exports	41.3	4.4	3.2	2.9	10.5	3.3	6.9	7.4	13.2
Processing exports	54.7	6.5	3.4	3.2	13.1	12.2	13.9	9.5	6.0
Other processing exports	4.0	0.1	0.2	0.2	0.5	0.9	0.6	0.4	1.7

Source: Updated from Rahman and Thorbecke (2007) and China (2006).

Notes: Hong Kong is included in China's imports from East Asia since inbound imports from Hong Kong are largely from other East Asian economies, generally intended for further processing into finished exports in China. These finished processing exports are then transshipped not only to the U.S. and EU-15, but also to the rest of Asia. However China's exports via Hong Kong are generally destined for the U.S. and EU-15 markets. EU-15 includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and United Kingdom. ASEAN-5 includes Indonesia, Malaysia, Philippines, Singapore & Thailand.

Feenstra and Spencer (2005, p.1) noted that both the "processing exports by foreign-owned firms" and "other processing exports by Chinese-owned firms" were largely produced under contractual arrangements with foreign multinationals, whereas the "ordinary exports by local firms" did not have these arrangements.

Table-2 (Contd.)

Panel-C: China's Trade Account Balance – 1993 and 2005 (in billions of U.S. dollars)

Partner Trade categories	World	Japan (1)	S. Korea & Taiwan (2)	ASEAN-5 (3)	East Asia (4=1+2+3)	Hong Kong	United States	EU-15	U.S.+ EU-15	Rest of the World
1993										
Trade account balance	-12.2	-7.5	-14.0	-1.3	-22.8	11.6	6.3	-3.5	14.4	-3.8
Ordinary trade	5.2	0.7	0.3	-0.1	0.9	7.7	0.0	-2.0	5.7	-1.5
Processing trade	7.9	-1.3	-9.4	-0.6	-11.4	5.7	9.7	4.2	19.5	-0.3
Other processing trade	-25.2	-6.9	-4.9	-0.6	-12.4	-1.7	-3.4	-5.8	-10.8	-2.0
2005										
Trade account balance	102.0	-16.4	-99.8	-23.8	**-140.1**	112.3	114.3	61.4	**287.9**	-45.8
Ordinary trade	35.4	-2.5	-12.9	2.0	**-13.4**	21.6	26.9	14.4	**62.9**	-14.0
Processing trade	142.5	4.5	-69.3	-13.3	**-78.1**	85.1	92.9	60.4	**238.4**	-17.9
Other processing trade	-75.9	-18.5	-17.7	-12.4	**-48.6**	5.6	-5.6	-13.4	**-13.4**	-13.9

Source: Updated from Rahman and Thorbecke (2007) and China (2006).

Notes: Following Kwan (2002), China's bilateral trade surplus against Hong Kong is considered as China's bilateral trade surplus against the U.S. and EU-15. Trade account balance regarding the "processing trade" is related to foreign affiliates of multinationals, while trade account balances regarding the "ordinary exports" and the "other processing exports" are related to Chinese-owned local firms (Feenstra and Spencer, 2005).

Table 3—**Estimation of Autoregressive and Distributed Lag Model for China's Processing Exports to 33 Countries, 1992-2005**
(Base Case Results for the True Model)

	1	2	3	4
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2
y_{it-1}	0.988***	0.774***	0.791***	0.776***
	(0.073)	(0.084)	(0.069)	(0.073)
y_{it-2}	-0.020	-0.072	0.038	0.048
	(0.070)	(0.059)	(0.051)	(0.058)
GDP_{it}	2.466***	2.480***	2.674***	2.560***
	(0.487)	(0.484)	(0.556)	(0.508)
GDP_{it-1}	-3.037***	-2.382***	-2.515***	-2.091**
	(0.770)	(0.739)	(0.715)	(0.797)
GDP_{it-2}	0.591	0.646	0.014	-0.297
	(0.452)	(0.451)	(0.392)	(0.521)
RER_{cit}	-0.784***	-0.799***	-0.718***	-0.754***
	(0.191)	(0.187)	(0.186)	(0.182)
RER_{cit-1}	0.523**	0.380	0.419*	0.429
	(0.252)	(0.231)	(0.246)	(0.263)
RER_{cit-2}	0.267*	0.082	0.254**	0.160
	(0.157)	(0.155)	(0.115)	(0.165)
RER_{wt}	-1.691***	-1.368***	-1.088***	-1.306***
	(0.436)	(0.270)	(0.224)	(0.368)
RER_{wt-1}	0.770**	0.763**	0.598*	0.533
	(0.351)	(0.340)	(0.295)	(0.340)
RER_{wt-2}	-0.799**	-0.738***	-0.439*	-0.715*
	(0.324)	(0.249)	(0.217)	(0.352)
m1			-2.94***	-2.91***
m2			-0.29	-0.28
Hansen J Statistic (d.f.)			0.009	0.590
			(10)	(28)
No. of Groups	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396

*Significance tests: *** p<0.01, ** p<0.05, * p<0.1*

Notes:

1. The estimated autoregressive and distributed lag (ADL) model (Equation 3.1):

$$y_{it} = \sum_{k=1}^{2}\alpha_k y_{it-k} + \beta_0 GDP_{it} + \beta_1 GDP_{it-1} + \beta_2 GDP_{it-2} + \xi_0 RER_{cit} + \xi_1 RER_{cit-1} + \xi_2 RER_{cit-2} +$$
$$\psi_0 RER_{wt} + \psi_1 RER_{wt-1} + \psi_2 RER_{wt-2} + \boldsymbol{\gamma}'\mathbf{z_i} + \eta_i + \boldsymbol{\delta}_i'\mathbf{d_{it}} + u_{it}$$

2. Hong Kong Export Price Index is used to deflate the nominal dollar value of China's processing exports.
3. Both GMM1 and GMM2 are one-step system GMM estimates. While GMM1 assumes that RER_{ci} and RER_W are exogenous, GMM2 assumes that they are predetermined. Appendix 2 provides more specific details on the moment conditions and the resulting extended instrument matrices that are used to derives the GMM estimates.
4. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses.
5. m1 and m2 are tests for first-order and second-order serial correlation in the first-differenced residuals, asymptotically distributed as N(0,1) under the null of no serial correlation.
6. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. P-value is reported.

Table 4— Estimation of Autoregressive and Distributed Lag Model for China's Ordinary Exports to 33 Countries, 1992-2005
(Base Case Results for the True Model)

	1	2	3	4
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2
y_{it-1}	0.657***	0.504***	0.572***	0.537***
	(0.104)	(0.096)	(0.078)	(0.105)
y_{it-2}	0.288***	0.198**	0.342***	0.384***
	(0.096)	(0.087)	(0.089)	(0.103)
GDP_{it}	1.795***	1.823***	1.768***	1.564**
	(0.466)	(0.487)	(0.628)	(0.635)
GDP_{it-1}	-2.185***	-1.796***	-1.703**	-1.369
	(0.718)	(0.674)	(0.793)	(0.943)
GDP_{it-2}	0.449	0.039	0.026	-0.120
	(0.422)	(0.404)	(0.289)	(0.358)
RER_{cit}	-0.861***	-1.001***	-0.821***	-0.889***
	(0.128)	(0.139)	(0.160)	(0.150)
RER_{cit-1}	0.526***	0.417***	0.500**	0.480***
	(0.154)	(0.147)	(0.188)	(0.169)
RER_{cit-2}	0.414***	0.233*	0.411***	0.352**
	(0.119)	(0.140)	(0.109)	(0.131)
RER_{wt}	-0.853***	-1.028***	-0.455**	-0.676**
	(0.315)	(0.246)	(0.214)	(0.328)
RER_{wt-1}	0.409*	0.348	0.436*	0.398*
	(0.225)	(0.231)	(0.224)	(0.225)
RER_{wt-2}	-0.904***	-1.119***	-0.701***	-0.937***
	(0.243)	(0.217)	(0.164)	(0.218)
ml			-1.71*	-1.84**
m2			-0.08	-0.34
Hansen J Statistic (d.f.)			0.009	0.289
			(10)	(28)
No. of Groups	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396

*Significance tests: *** p<0.01, ** p<0.05, * p<0.1*

Notes:

1. The estimated autoregressive and distributed lag (ADL) model (Equation 3.1):

$$y_{it} = \sum_{k=1}^{2} \alpha_k y_{it\text{-}k} + \beta_0 GDP_{it} + \beta_1 GDP_{it-1} + \beta_2 GDP_{it-2} + \xi_0 RER_{cit} + \xi_1 RER_{cit-1} + \xi_2 RER_{cit-2} +$$
$$\psi_0 RER_{wt} + \psi_1 RER_{wt-1} + \psi_2 RER_{wt-2} + \boldsymbol{\gamma}'\mathbf{z_i} + \eta_i + \boldsymbol{\delta}'_{\mathbf{i}}\mathbf{d_{it}} + u_{it}$$

2. Hong Kong Export Price Index is used to deflate the nominal dollar value of China's processing exports.
3. Both GMM1 and GMM2 are one-step system GMM estimates. While GMM1 assumes that RER_{ci} and RER_{W} are exogenous, GMM2 assumes that they are predetermined. Appendix 2 provides more specific details on the moment conditions and the resulting extended instrument matrices that are used to derives the GMM estimates.
4. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses.
5. ml and m2 are tests for first-order and second-order serial correlation in the first-differenced residuals, asymptotically distributed as N(0,1) under the null of no serial correlation.
6. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. P-value is reported.

Table 5— Estimation of Autoregressive and Distributed Lag Model for China's Processing Exports to 33 Countries, 1992-2005 (Base Case Results for the Hypothetical Model[1])

	1	2	3	4
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2
y_{it-1}	0.927***	0.805***	0.645***	0.663***
	(0.069)	(0.074)	(0.106)	(0.090)
y_{it-2}	0.039	0.027	0.196***	0.149**
	(0.067)	(0.060)	(0.071)	(0.060)
GDP_{it}	2.459***	3.133***	2.446***	2.494***
	(0.503)	(0.533)	(0.602)	(0.574)
GDP_{it-1}	-2.546***	-2.448***	-1.583**	-1.465**
	(0.804)	(0.813)	(0.752)	(0.704)
GDP_{it-2}	0.109	0.604	-0.705	-0.844*
	(0.495)	(0.497)	(0.454)	(0.482)
RER_{cit}	-0.484***	-0.432***	-0.579***	-0.584***
	(0.163)	(0.159)	(0.177)	(0.160)
RER_{cit-1}	0.145	0.069	0.220	0.178
	(0.176)	(0.164)	(0.169)	(0.167)
RER_{cit-2}	0.342***	0.374***	0.315***	0.236**
	(0.115)	(0.123)	(0.071)	(0.088)
m1			-2.36**	-2.81***
m2			-1.30	-1.11
Hansen J Statistic (d.f.)			0.012	0.111
			(10)	(21)
No. of Groups	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396

*Significance tests: *** p<0.01, ** p<0.05, * p<0.1*

Notes:
1. The estimated autoregressive and distributed lag (ADL) model (Equation 3.2):

$$y_{it} = \sum_{k=1}^{2} \alpha_k^* y_{it-k} + \beta_0^* GDP_{it} + \beta_1^* GDP_{it-1} + \beta_2^* GDP_{it-2} + \xi_0^* RER_{cit} + \xi_1^* RER_{cit-1} + \xi_2^* RER_{cit-2} + \boldsymbol{\gamma}^{*'} \mathbf{z_i} + \eta_i + \boldsymbol{\delta}_i^{*'} \mathbf{d_{it}} + v_{it}$$

The above model is misspecified as the variable RER_W is excluded and hence the production networks feature of China's processing exports is ignored.
2. Hong Kong Export Price Index is used to deflate the nominal dollar value of China's processing exports.
3. Both GMM1 and GMM2 are one-step system GMM estimates. While GMM1 assumes that RER_{ci} and RER_W are exogenous, GMM2 assumes that they are predetermined. Appendix 2 provides more specific details on the moment conditions and the resulting extended instrument matrices that are used to derives the GMM estimates.
4. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses.
5. ml and m2 are tests for first-order and second-order serial correlation in the first-differenced residuals, asymptotically distributed as N(0,1) under the null of no serial correlation.
6. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. P-value is reported.

Table 6— Estimation of Autoregressive and Distributed Lag Model for China's Ordinary Exports to 33 Countries, 1992-2005 (Base Case Results for the Hypothetical Model[1])

	1	2	3	4
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2
y_{it-1}	0.636***	0.629***	0.561***	0.545***
	(0.105)	(0.103)	(0.072)	(0.080)
y_{it-2}	0.301***	0.276***	0.287***	0.301**
	(0.097)	(0.098)	(0.096)	(0.111)
GDP_{it}	1.937***	2.544***	1.923***	1.850***
	(0.473)	(0.558)	(0.572)	(0.598)
GDP_{it-1}	-2.365***	-2.515***	-1.940**	-1.820**
	(0.687)	(0.726)	(0.714)	(0.741)
GDP_{it-2}	0.494	0.770*	0.180	0.130
	(0.401)	(0.414)	(0.274)	(0.315)
RER_{cit}	-0.728***	-0.639***	-0.717***	-0.749***
	(0.113)	(0.120)	(0.109)	(0.118)
RER_{cit-1}	0.243**	0.215*	0.241	0.204
	(0.121)	(0.120)	(0.146)	(0.142)
RER_{cit-2}	0.574***	0.663***	0.604***	0.559***
	(0.103)	(0.118)	(0.128)	(0.136)
m1			-1.66*	-1.68*
m2			0.14	0.06
Hansen J Statistic (d.f.)			0.005	0.087
			(10)	(21)
No. of Groups	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396

*Significance tests: *** p<0.01, ** p<0.05, * p<0.1*

Notes:

1. The estimated autoregressive and distributed lag (ADL) model (Equation 3.2):

$$y_{it} = \sum_{k=1}^{2} \alpha_k^* y_{it-k} + \beta_0^* GDP_{it} + \beta_1^* GDP_{it-1} + \beta_2^* GDP_{it-2} + \xi_0^* RER_{cit} + \xi_1^* RER_{cit-1} + \xi_2^* RER_{cit-2} + \boldsymbol{\gamma}^{*'} \mathbf{z_i} + \eta_i + \boldsymbol{\delta}_i^{*'} \mathbf{d_{it}} + v_{it}$$

The model is misspecified as the variable RER_W is excluded and hence the role of supply chain link of China's ordinary exports is ignored.

2. Hong Kong Export Price Index is used to deflate the nominal dollar value of China's processing exports.
3. Both GMM1 and GMM2 are one-step system GMM estimates. While GMM1 assumes that RER_{ci} and RER_W are exogenous, GMM2 assumes that they are predetermined. Appendix 2 provides more specific details on the moment conditions and the resulting extended instrument matrices that are used to derives the GMM estimates.
4. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses.
5. m1 and m2 are tests for first-order and second-order serial correlation in the first-differenced residuals, asymptotically distributed as N(0,1) under the null of no serial correlation.
6. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. P-value is reported.

Table A1

Panel Unit Root Tests: Levin-Lin-Chu ADF test statistics

Test Statistic Variables	Levin-Lin-Chu ADF t_δ test	Levin-Lin-Chu ADF t_δ^* test	Specifications for deterministic and/or autoregressive order in the error process
Real exports (ordinary)	-18.156***	-11.336***	Constant and trend; AR(1)
Real exports (processing)	-23.480***	-16.065***	Constant and trend; AR(1)
Real GDP (GDP_{it})	-12.990***	-6.897***	Constant and trend; AR(1)
RMB RER (RER_{cit})	-7.797***	-2.749***	Constant; AR(4)

'*', '**' and '***' denote 10%, 5% and 1% statistical significance, respectively.

Notes: RER_{cit} represents the bilateral real exchange rate of Chinese renminbi vis-à-vis country i. The weighted real exchange rate between China and East Asian countries that supply intermediate goods to China, RER_{wt}, is a time series variable. Since the variable RER_{wt} does not vary across cross-sections, we obtain both the augmented Dickey-Fuller and the Philips-Perron unit root test statistics. The p-values of the unit root test statistics are about 0.90. The results indicate that RER_{wt} is a unit root process, regardless of the number of higher-order autoregressive terms and/or a drift term included in the estimated regression.

Appendix-1

Persistency in the Individual DGPs

In order to assess the persistency of the individual variables, we estimate the univariate autoregressive model[1]

$$\Delta y_{it} = \alpha_i' d_{it} + \beta_i y_{it-1} + \sum_{L=1}^{P_i} \theta_{iL} \Delta y_{it-L} + e_{it} \ . \qquad \text{(A1)}$$

Here d_{it} is a vector of deterministic variables (e.g., intercept or time trend) and α_i is the corresponding vector of coefficients. Thus for the model without intercepts and trends, $d_{it} = \phi$ (the empty set); for the model with intercepts, $d_{it} = \{1\}$; and for the model with both intercepts and individual specific time trends, $d_{it} = \{1, t\}$. Here α_{1i} is assumed to represent cross-section specific intercepts capturing the unobserved fixed effect parameter η_i and e_{it} is assumed to have finite moments and in particular $E(e_{it}) = E(e_{is} e_{it}) = 0$, for $i = 1, \cdots, N$ and $\forall s \neq t$. For the real export and real GDP series, we estimate the model including both the deterministic variables (i.e., both the cross-section specific intercepts and trend term). For the RMB real exchange rate variable, we estimate the same autoregressive specification but without the trend element. By contrast, we derive the estimate of autoregressive parameter for the time series variable, RER_{wt}. The choice of appropriate autoregressive order and deterministic terms for all the cross-sectionally and time-varying series is based on the Levin-Lin-Chu unit root results that are presented in

[1] The form is the Sims, Stock and Watson (1990) canonical form for higher order autoregressive processes, originally proposed by Fuller (1976).

Table A1. In addition to the OLS and within estimates, we employ both GMM first-difference and GMM system estimators. GMM first-difference estimators are based on $m=(T-p-1)(T-p)/2$ linear moment conditions that are defined as $E(y_{it-s}\Delta e_{it})=0$ for $t=p+1,\cdots,T$ and $s\geq 2$. By contrast, GMM system estimators are based on an extended set of moment conditions, which additionally includes another $(t-p-1)$ linear moment conditions that are defined as $E(e_{it}\Delta y_{it-1})=0$ for $t=p+1,\cdots,T$.

Table A2
Estimates of the autoregressive parameter of individual data generation processes (DGPs)

Name of the DGPs	OLS	GMM-Sys	Within	GMM-Diff
Ordinary real exports ($EX1_{it}$)	0.978***	0.928***	0.719***	0.811***
	(0.013)	(0.043)	(0.063)	(0.091)
Processing real exports ($EX2_{it}$)	0.979***	0.912***	0.754***	0.360***
	(0.009)	(0.027)	(0.060)	(0.097)
Real GDP (GDP_{it})	0.997***	0.974***	0.771***	0.521***
	(0.001)	(0.008)	(0.043)	(0.157)
RMB RER (RER_{cit})	0.999***	0.786***	0.282***	-0.070
	(0.018)	(0.081)	(0.094)	(0.132)
Weighted RER (RER_{wt})	.881***	n/a	n/a	n/a
	(.091)			

Notes: *** p<0.01, ** p<0.05, * p<0.1

Notes: The data series $EX1_{it}$ and $EX2_{it}$ represent China's bilateral ordinary and processing exports respectively to country i. GDP_{it} represents real gross domestic product in the importing country. RER_{cit} represents the bilateral real exchange rate of Chinese renminbi vis-à-vis country i. RER_{wt} represents the weighted real exchange rate between China and countries providing imports for processing to China. Only the OLS estimate of persistency of RER_{wt} is reported, since the variable does not vary cross-sectionally.

Table A2 presents our results on persistency characteristics. The OLS and within estimates of the autoregressive parameter in a dynamic panel data model, which are biased upwards and downwards respectively, are reported for comparative purpose. GMM system estimators provide better estimates of the true parameter. The evidence indicates that, though the dependent variable (the ordinary exports or the processing exports) are both trend stationary series, they are highly persistent. GMM system estimate of the RMB RER coefficient also indicates to a high degree of persistency. Similarly the RER_{wt} variable is also found to be highly persistent. These results imply that GMM first-difference estimators in our multivariate dynamic panel model are likely to be weakly identified and hence inconsistent. Moreover if the mode allows for endogeneity in the real exchange rate variables, the weak instrument problem appears to be more plaguing in both Anderson-Hsiao IV estimators and Arrelano-Bond GMM first-difference estimators. The findings provide unequivocal evidence that the multivariate dynamic panel data model in the present case must resolve on the potential weak instrument problem. The study therefore exploits the extended instrument matrix as proposed by Blundell and Bond (1998) and derives relatively consistent and efficient GMM system estimators.

Appendix-2

Moment Conditions and the Instrument Matrix

This appendix provides details on the moment conditions and related GMM-style instrument matrices. For brevity, we do not discuss the time-invarying set of variables that are relevant IV-style instruments and have been used in actual GMM system estimations.

1. Moment Conditions and Instrument Matrix for GMM1 Estimators

The GMM1 estimation results are obtained by assuming that all of the right hand side variables are exogenous. The vector $\mathbf{x}'_{it} = (RER_{cit}\ \ RER_{wt}\ \ GDP_{it})$ represents the set of time-and cross-sectionally varying variables, the vector $\mathbf{z}_i$ represents the set of time-invarying variables. The variables are described in the text. The basic moment conditions in our ADL(2,2) model are defined below:

i. For the difference equations: $E(y_{it-s}\Delta u_{it}) = 0$ and $E(\Delta x'_{it}\Delta u_{it}) = 0$ for $t = 4,\cdots,T$ and $s \geq 2$; and

ii. For the level equations: $E(\Delta y_{it-1}u_{it}) = 0$ and $E(x'_{it}u_{it}) = 0$ for $t = 4,\cdots,T$.

The instrument matrix that arises from condition (i) can be written as:

$Z_i = \text{diag}(y_{i1}\ y_{i2}\cdots y_{is} \vdots \Delta\mathbf{x}'_i)$ for $t = 4,\cdots,T$ and $s \geq 2$, where $\Delta\mathbf{x}'_i = (\Delta\mathbf{x}'_{i4}\cdots\Delta\mathbf{x}'_{iT})$. This instrument matrix is used to obtain Arellano-Bond GMM first-difference estimators. Adding condition (ii) results in the extended instrument matrix as suggested by Blundell and Bond (1998):

$$Z_i^+ = \begin{bmatrix} Z_i & 0 & 0 & \cdots & 0 & \vdots & 0 \\ 0 & \Delta y_{i3} & 0 & \cdots & 0 & \vdots & \mathbf{x}'_{i4} \\ 0 & 0 & \Delta y_{i4} & \cdots & 0 & & \mathbf{x}'_{i5} \\ \vdots & \vdots & \vdots & \ddots & \vdots & \vdots & \vdots \\ 0 & 0 & 0 & \cdots & \Delta y_{i,T-1} & . & \mathbf{x}'_{iT} \end{bmatrix}$$

The subscript 3 in our model refers to 1994, the subscript 4 to 1995, and the subscript T to 2005. The reported GMM1 estimates are obtained by utilizing on the extended instrument matrix.

2. Moment Conditions and Instrument Matrix for GMM2 Estimators

The GMM2 estimation results are obtained by assuming that both the RER variables are predetermined and that the other right-hand side variables are exogenous. Let $\mathbf{x}_{1,it}$ represents the set of predetermined variables and $\mathbf{x}_{2,it}$ represents the exogenous set of variables. Here $\mathbf{x}'_{1,it} = (RER_{citt}\ \ RER_{wt})$ and $\mathbf{x}'_{2,it} = (GDP_{it}\ \ \mathbf{z}'_i)$. The variables are described in the text. The moment conditions for the ADL(2,2) are defined below:

i. For the difference equations: $E(y_{it-s}\Delta u_{it}) = 0$, $E(\mathbf{x}'_{1,it-s+1}\Delta u_{it}) = 0$ and $E(\Delta x'_{2,i}\Delta u_{it}) = 0$ for $t = 4,\cdots,T$ and $s \geq 2$, where $\mathbf{x}'_{2,i} = (\mathbf{x}'_{2,i4}\cdots\mathbf{x}'_{2,iT})$ and

ii. For the level equations: $E(\Delta y_{it-1}u_{it})=0$, $E(\Delta \mathbf{x}'_{1,\mathbf{it}}u_{it})=0$, and $E(x'_{2,it}u_{it})=0$ for $t=4,\cdots,T$.

The instrument matrix that arises from the condition (i) is compactly written as $Z_i = \text{diag}(y_{i1}\ y_{i2} \cdots y_{is}\mathbf{x}'_{1,i1}\cdots \mathbf{x}'_{1,is+1} \vdots \Delta\mathbf{x}'_{2,i})$ for $t=4,\cdots,T$ and $s \geq 2$, where $\Delta\mathbf{x}'_{2,i} = (\Delta\mathbf{x}'_{2,i4}\cdots\Delta\mathbf{x}'_{2,iT})$.

This instrument matrix is used to obtain Arellano and Bond GMM first difference estimators. By contrast, combining both the moment conditions (i) and (ii) results in the extended instrument matrix as suggested by Blundell and Bond (1998). The extended instrument matrix is shown below:

$$Z_i^+ = \begin{bmatrix} Z_i & 0 & 0 & \cdots & 0 & \vdots\ 0 \\ 0 & \Delta y_{i3}\ \Delta\mathbf{x}'_{1,i4} & 0 & \cdots & 0 & \vdots\ \mathbf{x}'_{2,i4} \\ 0 & 0 & \Delta y_{i4}\ \Delta\mathbf{x}'_{1,i5} & \cdots & 0 & \mathbf{x}'_{2,i5} \\ \vdots & \vdots & \vdots & \ddots & \vdots & \vdots\ \vdots \\ 0 & 0 & 0 & \cdots & \Delta y_{i,T-1}\ \Delta\mathbf{x}'_{1,iT} & .\ \mathbf{x}'_{2,iT} \end{bmatrix}$$

The subscript 3 again refers to 1994, the subscript 4 to 1995, and the subscript T to 2005.

In both cases of GMM1 and GMM2 estimation results, we employ the collapsed version of this instrument matrix to avoid the problem of overfitting the instrumented variable. As an alternative to collapsing the instruments, we also set lag limits to select fewer than available instruments. The results are invariably similar. Furthermore, in both GMM1 and GMM2 cases, we include the set of time-invarying gravity variables and trend terms as the IV-type instruments for the level equations. This is intended to minimize the problem of potential panel heterogeneity in our observed data set.

3. Moment Conditions and Instrument Matrix for GMM3 Estimators

The GMM3 estimation results are obtained by assuming that both the RER variables are endogenous and that the other right-hand side variables are exogenous. Assuming that the vector $\mathbf{x}_{1,it}$ represents the set of endogenous variables and $\mathbf{x}_{2,it}$ represents the exogenous set of variables. the moment conditions for the ADL(2,2) are then as follows:

i. For the difference equations: $E(y_{it-s}\Delta u_{it})=0$, $E(\mathbf{x}'_{1,\mathbf{it-s}}\Delta u_{it})=0$ and $E(\Delta x'_{2,i}\Delta u_{it})=0$ for $t=4,\cdots,T$ and $s\geq 2$, where $\mathbf{x}'_{2,i} = (\mathbf{x}'_{2,i4}\cdots\mathbf{x}'_{2,iT})$ and

ii. For the level equations: $E(\Delta y_{it-1}u_{it})=0$, $E(\Delta\mathbf{x}'_{1,\mathbf{it-1}}u_{it})=0$, and $E(x'_{2,it}u_{it})=0$ for $t=4,\cdots,T$.

Note that potential endogeneity of a variable x_{it} requires that its lags dated $(t-2)$ or earlier can only be the instruments in dynamic panel model that includes unobserved fixed effects. If x_{it} is in fact not endogenous, the assumption is restrictive and causes weak instrument problem. The GMM3 estimates that are based on the above moment conditions are found to be weakly identified and further downward biased than the FE estimates. We therefore do not report them.

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