THE IMPACT OF A COMMON CURRENCY ON EAST ASIAN PRODUCTION NETWORKS AND CHINA'S EXPORTS BEHAVIOR [*]

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Abstract

Vertical fragmentation of product value chain across borders is the driving force of growing economic interdependency in East Asia. A common currency, not flexible exchange rates between national currencies, would reduce flexibility in relative prices within East Asia. Its impact would be far greater for exports that have stronger production network linkage. In order to test the hypothesis, the paper estimates the effect of a common currency on China's processing and ordinary exports separately. The distinction is necessary because the processing exports, unlike the ordinary exports, are produced along the regional production networks, with final stages of assembly and exporting being increasingly concentrated in China. The short-run dynamics indicate that the effect on China's processing exports is more than double the corresponding effect on China's ordinary exports. The long-run effect on the processing exports of intra-regional RER flexibility, which is otherwise the lack of a regional currency, is almost nine times as large as the long-run effect of a unilateral RMB appreciation. By contrast, the corresponding long-run effect is statistically insignificant for the case of ordinary exports that are produced primarily by using local inputs. The long-run coefficient of this intra-regional RER flexibility implies that the actual volume of processing exports is 20 percent below the potential. The magnitudes of these effects are consistent with the hypothesis that a common currency would further integrate East Asian production networks and promote regional economic integration.

Keywords: Production networks, fragmentation of value chain, optimum currency area, common currency, exchange rate flexibility, China.
JEL classification: F14, F33, F36, F42

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1. Introduction

In a seminal paper, Robert Mundell (1961) argued that an optimum currency area (OCA) would be a region, not the domain of national currencies. His argument is that if factors are sufficiently mobile across national boundaries in the region, then a flexible exchange rate system based on national currencies becomes unnecessary, and may even be positively harmful. McKinnon (1963) further advanced the concept in terms of the ratio of tradable to non-tradable goods. His argument is that if a number of countries trade extensively with each other and if each pegs its currency to a representative bundle of imports, then each currency will be pegged to the others. McKinnon argued that to maintain the liquidity value of individual currencies, a fixed exchange rate system, or a common currency, would be necessary. A common currency would greatly facilitate contractual arrangements. It would thereby stimulate factor mobility among the countries and promote economic specialization and growth in the region. In effect, both the OCA criteria are conceptually interrelated and endogenous to the intra-regional trade integration. If countries in a particular region are increasingly integrated in their production and trading relationships, a new demand will arise for a regional currency and against national currencies within the region.

The recent literature generally suggests that capital and technology have become highly mobile across East Asian countries since the early 1990s[1]. The literature further indicates that East Asian economies have synchronized business cycles, particularly after the Asian financial crisis and that there is a greater degree of interdependency among the countries in the region. In the context of East Asian production networks, it is rational to argue that the condition of greater labor mobility is largely attained by vertical fragmentation of production processes across borders in East Asia[2]. Note that the notion of labor mobility is meant not to be in terms of geographical and/or inter-industry dimensions as indicated in the OCA literature, but by the way of intra-industry fragmentation of product value chain across national borders. The present study therefore conjectures that East Asia is an optimum currency area. However, East Asian countries have their independent national currencies and pursue heterogeneous exchange rate policies. This leads to the research question: what would happen to East Asian production networks and regional trade integration, had there been a common currency? To put it differently, what is the opportunity cost for East Asian exports for not having a regional currency? Particularly, this paper's focus is on measuring the costs to China's processing and ordinary exports respectively.

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[1] See, for example, Kwan (2001), Hatch, (2003), McKinnon and Schnabl (2003), Yusuf et al. (2004), Thorbecke and Yoshitomi (2006), Fujita (2007), Kawai (2007), and Plummer and Wignaraja (2007). The author thanks Ronald McKinnon for referring to recent empirical evidence on increasing synchronization of business cycles in East Asia.

[2] The phenomenon of international fragmentation of production processes can be construed much in line with Raymond Vernon's (1966) product cycle hypothesis. Vernon's fundamental conjecture that the locus of production will be shifted to the less-developed South as the production techniques become standardized implicitly recognizes that parts of production process, not the entire product value chain, can also be shifted to the less-developed South. The resultant trade pattern would be vertical intra-industry trade, not the horizontal inter-industry trade.

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[3]. Anderson and van Wincoop (2003), assuming complete production specialization and homothetic preferences, obtained a theoretical gravity equation. The model predicts that bilateral trade, after controlling for size, depends on the bilateral trade barrier between two regions, relative to the product of their multilateral resistance indices[4]. They estimated the model both in the context of a two-country setting consisting of the U.S. states and Canadian provinces, and a multi-country setting that also included 20 other industrialized countries. Based on the estimate of elasticity of substitution (i.e., 5.0) from Hummels (1999), the study found that a tariff equivalent estimate of the U.S.-Canada border barrier would be 48 percent. The study further found that border barrier reduced trade between the United Sates and Canada by 44 percent of that of border-less trade. Rose and van Wincoop (2001), using 1980 and 1990 data for a set of 143 countries, estimated the Anderson-van Wincoop gravity model to estimate the effect on trade of monetary unions. The study replaced the multilateral resistance terms with country-specific fixed effects. They found that the tariff equivalent estimate of the monetary barrier to trade would be 26 percent. For the case of European Monetary Union (EMU), the results showed a 58 percent trade-creating effect of currency union for the euroland countries. This is perhaps the most conservative estimate of the effect of a currency union on trade. Klein and Shambaugh (2006) developed a comprehensive database for 181 countries over the 1973-1999 period and used a *de facto* exchange rate regime classifications. The *de facto* classification scheme included currency union, direct peg and indirect peg, all being mutually exclusive meaning that any one observation can only be coded as one type of exchange rate regime. They also estimated a standard gravity model including dummies for three regime classifications and other usual controls. The study found that a fixed exchange rate system (direct peg) would 'increase international trade with one another 36 percent relative to intranational trade (p. 370).'

Surprisingly, all the empirical results are, to the best of our knowledge, *ex post* estimates of the effect of either a common currency or a fixed exchange rate system[5]. The applied methodologies, in general, do not suggest a way to obtain *ex ante* estimates of the effect of a common currency, when a region is an optimum currency area but with independent national currencies linked by flexible exchange rates. Nor does the literature offer a

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[3] Rose and Engel (2002) and Frankel and Rose (2002) showed that a currency union would increase trade between union members by a factor of over three. Rose and Stanley (2005, p23) reported 34 empirical estimates of the effect on trade of currency union. The median coefficient of the currency union dummy was shown to be 1.2, implying that membership in a currency union would, *cateris paribus*, triple bilateral trade—$e^{1.20}>3.0$.

[4] See Kalirajan (2007) for a stochastic frontier formulation of standard gravity model to estimate the effect of country-specific resistance on bilateral trade flows.

[5] Klein and Shambaugh (2006) showed that the number of observations with *de facto* fixed exchange rate systems including currency union in a typical database of bilateral trade and exchange rates would be about 2 percent of total observations. They represent only 11.5 percent of world trade. A dummy variable representation for those limited observations in a large dataset is non-random and likely to draw misleading information into the variance-covariance matrix. Estimation inefficiency can be substantial depending on *de facto* regime classifications and empirical specifications.

framework to estimate the effect for a particular country, which belongs to a currency area. The present study develops a conceptual framework to that end and obtains empirical estimates of the effect of a common currency in East Asia.

The conceptual framework essentially incorporates the features of cross-border fragmentation of production processes and intra-regional exchange rate flexibility into the modeling. Since export production has evolved regionally with the value chain being fragmented vertically across national boundaries, the "gross value" of exports from any East Asian country represents the sum of incremental value-added that occurs along the cross-border production networks. Bilateral trade flows, which are recorded at "gross," within the region are thus mostly vertical intra-industry trade (VIIT). This causes the effect of intra-regional exchange rate flexibility on final exports to be multiplicative by the degree of fragmentation of value chain of those exports. The study uses alternative proxies to represent production network linkage of China's final exports with other East Asian countries. An inner product of the degree of production network linkage and the log of bilateral real exchange rate between China and China's regional trading partners thus constitute the key variable. The variable is called the intra-regional real exchange rate (RER) flexibility, which represents the unintended misalignment in relative prices between China and the rest of East Asia. The variable along with other relevant covariates enters into a multivariate modeling to explain bilateral real exports in an imperfect substitutes framework. The other covariates include bilateral RER variable, GDPs of importer and exporter, proxy for supply shift effect, and general gravity variables. This model is termed as fully specified model and the parameter estimate of the key variable of this model denotes the impact of intra-regional RER flexibility on exports. Another hypothetical model is then estimated, assuming that East Asia is an optimum currency area and hence that the intra-regional RER flexibility variable is irrelevant. That is, the hypothetical model includes all the controls but the RER flexibility variable. The estimate of bilateral RER variable in the hypothetical model and that of the intra-regional RER flexibility variable in the fully specified model are then statistically reconciled in order to estimate the impact of a common currency on exports. A testable hypothesis is that a common currency would have relatively larger effect on those exports that have deeper production network linkage across borders in East Asia.

The study applies the above framework to estimate the effect of a common currency on China's exports behavior. In doing that, the study distinguishes between China's processing exports and the ordinary exports. The distinction is imperative because the processing exports, unlike the ordinary exports, are produced along the regional production networks, with final stages of assembly and exporting being increasingly concentrated in China. As an outcome, China experiences increasing vertical intra-industry trade (VIIT) with the rest of East Asia and a surge in final exports to the United States, Europe and elsewhere. A common currency in East Asia would significantly eliminate flexibility in relative prices between China and the rest of East Asia. Its impact would therefore be far greater for the processing exports than it would be for the ordinary exports.

The study formulates China's export demand equation in an imperfect substitutes framework. An autoregressive and distributed lag (ADL) specification of the model is estimated for both the panels of China's processing and ordinary exports. The paper does not make the arbitrary assumptions that the variables are unit root processes and that there exist cointegration relations. Instead, exact time series properties of the data are obtained. The ADL specification is consistently estimated so that spurious estimates of the long-run parameters are not obtained. This is very likely when there are no cointegration relations in the observed data, but an arbitrary long-run model is estimated. Since the dynamic model includes fixed effects and one or more of the right-hand side variables can be predetermined and/or endogenous, both the pooled OLS and covariance estimators are inconsistent. Thus, the model is estimated by using Generalized Method of Moments (GMM) approach as suggested by Holtz-Eakin et al. (1988), Arellano and Bond (1991) and Blundell and Bond (1998). The consistent GMM estimators are based on a set of moment conditions that are related to both the differenced equations and the levels equations of the model.

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 it's 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 discusses time-series properties of the observed data. Section 4.2 introduces the dynamic panel data model. 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

Three economic regions of the world, i.e., NAFTA (North American Free Trade Area), EU-15 (European Union-15), and East Asia-15, registered rapid growth relative to the world average over the 1985-2005 period. Figure 1 shows the pattern of growth and economic interdependency of these three economic regions. The question is whether the economic interdependency within each region has been growing stronger in relation to the growth. The figure shows that it has been so in the case of East Asia[6]. In fact, the share of intra-regional trade of East Asia is approaching to that of EU. The most distinctive feature of East Asian growth and trade integration is that production is organized internationally in the region. The phenomenon is widely known as East Asian production networks.

[Insert Figure 1 here]

The production network underlies fragmentation of product value chain across borders in East Asia[7]. Here Japan and NIEs-2 (South Korea and Taiwan), in general, organize those

[6] Unlike NAFTA and EU-15 that shares stronger political and/or monetary union, East Asian economic integration has been mainly through market mechanisms, with little support from region-wide political institutions (Fujita, 2007).

[7] IBM's CEO, Sam Palmisano (2006), called it 'globally integrated business strategy' of multinational corporations. He argued that the global integration of production was not just to cut costs, but more to tap

production processes that use relatively higher skilled workers and produce sophisticated product prototypes, high-tech intermediate goods and capital equipments. These intermediate goods are transformed into finished products at assembly plants mainly in China. The finished products are then exported throughout the world. Palmisano (2006) noted that an estimated 60,000 manufacturing plants were built by foreign firms in China alone between 2000 and 2003 and that most of these factories target the global market, not the local Chinese market. Greenspan (2005) thus argued, "...production within Asia has evolved, with the final stages of assembly and exporting becoming increasingly concentrated in China." As an outcome, trade along these production networks, which is called vertical intra-industry trade (VIIT), has increased substantially over time. Figure 2 shows the surging pattern of intra-industry trade between China and the rest of East Asia. Its implication is 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 finished exports to the Americas, Europe and elsewhere.

[Insert Figure 2 here]

Table-1 reflects 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[8]. The export statistics are compiled into 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 exports* by Chinese owned firms. Similar nomenclature is followed for the compilation of import statistics. 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, whereas the ordinary exports did not have these arrangements.

Panel A of Table-1 shows that in 2005, the processing imports that are made under contractual arrangements with foreign multinationals accounted for 58 percent of China's overall imports. Of this 58 percent, about 65 percent came from other East Asian countries. By contrast, only 13 percent came from the U.S. and EU-15. Panel B of Table-1 shows that in 2005, China's processing exports accounted for 55 percent of its overall exports. Of this 55 percent, about 76 percent went to the Americas and Europe, if Hong

new sources of skills and knowledge. The strategic decisions are not simply a matter of off-loading noncore activities, nor are they mere labor arbitrage. They are about actively managing different operations, expertise, and capabilities across national boundaries. Fujita (2007, p. 18) argued that it 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.

[8] These data were made available to the author by RIETI, which purchased the data from China's Customs Statistics Information Center, Economic Information Agency, Hong Kong.

Kong is arguably treated as an entrepôt of trade and transshipment to the west[9]. Because processing imports are brought-in duty-free primarily for using in the production of finished exports, the dollar cost of these imports represents a substantial share of the gross value of China's processing exports. But, since exports by country are recorded on a gross basis rather than as value added, the widening bilateral deficits of both the U.S. and Europe against China, measured gross, have largely been in lieu of their wider deficits with other East Asian economies. Greenspan (2005) also emphasized on this point in his testimony before the U.S. Senate Finance Committee.

Panel C of Table-1 shows China's trade account balance in 1993 and 2005. In 2005, China incurred a deficit of $140 billions against the rest of East Asia, but a surplus of about $290 billions against the U.S. and EU-15. In terms of product types, almost 90 percent of China's bilateral trade deficit against the rest of East Asia occurred on account of processing trade (both [b] & [c]). By contrast, about 83 percent of China's bilateral trade surplus against the U.S. and EU-15 occurred on account of the processing trade. However, it is trade in intermediate inputs that define China's trade deficit with the rest of Asia, whereas it is trade in final goods that define China's trade surplus with the rest of the world. Therefore, it is the production and exporting of processed goods that are defining parameters of China's integration backward with East Asian production networks and forward into the world trading system.

The present study thus distinguishes between the processing exports and ordinary exports for analyzing the impact of a common currency 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. In order to do it, the study first aims to provide consistent estimates on the impact of RER flexibility between China and other East Asian countries that supply intermediate goods to China. The study then estimates potential costs respectively for the processing and the ordinary exports for not having a common currency in East Asia. Accordingly, a conceptual framework is developed in the following section.

3. Conceptual Framework of the Study

As evident in the previous section, the 'gross value' of China's exports essentially 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. Figure 3 depicts a schematic view of the fragmented value chain and real exchange rate relationships within East Asia and between East Asia and the rest of the world. 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 Asia. Since bilateral

[9] 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).

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 the incremental x_C, which is the Chinese value added. 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 (renminbi) exchange rate in the network. Though the vertical fragmentation 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 how changes in country i's nominal exchange rate, say the quasi-global currency like the U.S. dollar or its new rival the Euro, would enter into the production networks and affect VIIT as well as final exports from China. The related hypothetical question is what would happen, had there been a common currency in East Asia.

[Insert Figure 3 here]

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 (see, e.g., 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 Fig. 3, 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} [10]. Ogawa and Yang (2006; p. 17) find that East Asian countries do not have any effective coordination mechanism in

[10] 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. The time subscript t is suppressed for notational convenience.

their international macroeconomic policies and that their exchange rate policies are largely asymmetric to movements in the world invoice currencies. They even resort to competitive devaluation. 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 the intra-regional real exchange rates. In the context of China, the asymmetric reactions imply substantial variability in RER_{jc}. In other words, price stability along the production networks will be distorted, thereby affecting growing supply-chain linkages between China and the rest of East Asia. Its impact on China's final exports would be compounded by the degree of supply chain linkage of those final exports with the rest of East Asia. How do we capture the impact of intra-regional exchange rate flexibility on exports from China? The present study thus creates a new variable, called the intra-regional RER flexibility variable, which is defined below.

Let ω_j be the weight of country j in th*e* gross value of Chinese final exports. Then, the term $RER_{jc}^{\omega_j}$, which is $[\omega_j \ln(RER_{jc})]$ after log transformation, captures both the dynamic integration of China with country j and the asymmetry in exchange rate and monetary policies between them. For the present analysis, the new variable is thus defined as $RER_w = \Sigma_j \omega_j RER_{jc}$. The term ω_j is proxied either by the share of country j in China's imports for processing or by their intra-industry trade intensity. 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 RER flexibility between East Asian countries that organize larger part of the fragmented (cross-border) value chain and China where final stages of assembly are done to produce "processing exports." Its flexibility is the unintended misalignment in relative prices between China and the rest of East Asia[11]. 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

[11] Since $RER_{jc} \equiv (RER_{ji} - RER_{ci})$, all being in natural logs, we find that $RER_w = [\Sigma_j \omega_j RER_{jc}] \equiv [(\Sigma_j \omega_j RER_{ji}) - \{(\Sigma_j \omega_j) RER_{ci}\}]$. Let $s_t^* = RER_w$, $s_{1,i} = (\Sigma_j \omega_j RER_{ji})$ and $s_{2,i} = \{(\Sigma_j \omega_j) RER_{ci}\}$, the null of real exchange rate parity between China and the rest of East Asia holds if the model $s_t^* = b_1 s_{1,it} + b_2 s_{2,it} + u_{it}$ is stationary and $\mathbf{b}' = (b_1, b_2) = (1, -1)$. Here, $s_{1,it}$ represents the weighted real exchange rate between all the East Asian countries other than China and country i and $s_{2,it}$ the bilateral real exchange rate between China and country i. The subscript i indexes the panel of China's bilateral trading partners. The rejection of the null $(b_1 + b_2) = 0$ in favor of the alternative $(b_1 + b_2) \neq 0$ would imply significant flexibility in RER_w. Based on a dynamic panel estimation, we find the evidence of significant real misalignment that $\hat{b}_1 = 0.75$, $\hat{b}_2 = -0.98$, and the linear combination of $(\hat{b}_1 + \hat{b}_2) = -0.225$, all being statistically significant at 1%.

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[12]

$$y_i = \beta_1 RER_w + \beta_2 RER_{ci} + u_i. \tag{1}$$

Here β_1 measures the impact of the RER flexibility 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 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.

Let the hypothetical model, which ignores the influence of VIIT and the existence of asymmetry in exchange rate management in East Asia be

$$y_i = \beta_2^* RER_{ci} + v_i. \tag{2}$$

In this case, the RER_w variable is excluded. However, the 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 'fully specified model' from econometric estimation 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} (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. It implies that RER_{ci} is the relevant variable but RER_w is the irrelevant variable in explaining China's exports behavior. Hence, the hypothetical model

$$y_i = \beta_2 RER_{ci} + v_i \tag{2a}$$

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_w would cause $\beta_1 \to 0$.

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

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. The coefficient β_1 in the fully specified model (1) would measure the impact of the RER flexibility between China and the rest of East Asia. But it has indirect effect too. Note that the RMB coefficient in model (1) is $\beta_2 = (\hat{\beta}_2^* - \beta_1 b_{12})$, where $\hat{\beta}_2^* < 0$ and $\beta_1 b_{12} > 0$. In other words, the RMB coefficient β_2 is inflated by an absolute $\beta_1 b_{12}$ term. Thus, the cost to China's exports for not having a common currency would be measured by $(\beta_1 + \beta_1 b_{12})$[13]. In other words, had there been a common currency in East Asia and hence stable relative price relationships within East Asian countries, the coefficient $\beta_1 \rightarrow 0$ and hence "the enhanced effect" $\beta_1 b_{12} \rightarrow 0$. This means that $(\beta_1 + \beta_1 b_{12}) \rightarrow 0$. The hypothesis is that $\left|\beta_1 + \beta_1 b_{12}\right|_{PX} \succ \left|\beta_1 + \beta_1 b_{12}\right|_{OX}$, where the subscripts PX and OX denote the processing exports and the ordinary exports respectively.

It is of relevance to know the extent of loss in potential trade due to the presence of intra-regional RER flexibility. In the empirical gravity literature, trade effect of a currency union is estimated by $exp(\hat{\phi})$, with $\hat{\phi}$ being the coefficient of the indicator variable that is unity if two countries share a common currency, and zero otherwise. This is, however, inapplicable in the present case because our estimate of the effect of common currency is not based on the use of a dichotomous variable. It is rather the non-linear combination $\phi^* = (\beta_1 + \beta_1 b_{12})$ measuring the effect of RER_w, which is otherwise a time-varying common forcing variable in the model. Though RER_w captures the effect of intra-regional RER flexibility, the variable itself is not a measure of risk factor. Following Hooper and Kohlhagen (1978, p.500), the present study takes the absolute difference between RER_w and its fitted values based on a log linear trend equation to be the indicator of real exchange risk. Note that this exchange risk arises only from the variability of Chinese real exchange rates against other East Asian countries, not all the bilateral trading partners. It is defined as $d_t = \left|RER_{wt} - R\hat{E}R_{wt}\right|$, with $R\hat{E}R_{wt}$ being the linear prediction of RER_{wt} obtained from a log-linear trend equation[14].

The study then creates an adjustment term defined as $exp(d_t)^{\hat{\phi}^*}$. The adjustment term has temporal variation but uniform across cross-sections. If either $d_t \rightarrow 0$ or $\hat{\phi}^* \rightarrow 0$, the

[13] If the auxiliary coefficient b_{12} tends to be zero, the effect of a common currency would be measured by the coefficient β_1 per se.

[14] The definition of d_t assumes implicitly that a common currency arrangement would establish a stable but trending relative price relationship between East Asian countries. It thereby precludes the assumption that the relative prices be fixed. Hooper and Kohlhagen (1978) argued that the major advantage of this measure of risk, compared with the standard deviation measures obtained from either a log-linear trend equation or a first-order autoregressive equation, was that under pegged but adjustable exchange rates it might better indicate the market's assessment of exchange risk. Kenen and Rodrik (1986) used alternative standard deviation measures for estimating trade effect of short-term volatility in real exchange rates. This is however not the purpose of the present study.

term $exp(d_t)^{\hat{\phi}^*} \to 1$, otherwise $exp(d_t)^{\hat{\phi}^*} < 1$ because the coefficient $\hat{\phi}^* \leq 0$ by assumption. We can now define the potential trade $y^*_{it} = \hat{y}_{it} / exp(d_t)^{\hat{\phi}^*}$, where $\hat{y}_{it}$ is the predicted value of Chinese exports by estimating the fully specified model. If the term $exp(d_t)^{\hat{\phi}^*}$ is unity, $\hat{y}_{it}$ itself is the potential exports. On the other hand, if the adjustment term $exp(d_t)^{\hat{\phi}^*}$ is less than unity, the potential exports $y^*_{it} = \hat{y}_{it} /[exp(d_t)^{\hat{\phi}^*}]$ shall exceed the actual exports. Actual trade relative to the potential trade (y_{it} / y^*_{it}) is the intended measure of trade effect due to intra-regional RER flexibility between China and the rest of East Asia. For each cross-section i, the study reports both $avg(y^*_{it} - y_{it})$ and $(\bar{y}_i / \bar{y}^*_i)$ where the over-bar indicates the average taken over the time period.

This now clarifies how greater flexibility in East Asian real exchange rates that arise from independent national currencies and heterogeneous exchange rate policies would affect East Asian production networks and, therefore, those exports that have stronger production network linkage. 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 Time Series Properties of the Data

Conventional practice has been to assume that the observed data $(\mathrm{y}_{\mathrm{it}}, \mathbf{x}_{\mathbf{it}})$ 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.

In order to know whether non-stationarity in the data is due to a deterministic time trend or unit root, the study conducts panel unit root tests for the main data generation processes (DGPs). They include three cross-section and time series, i.e., bilateral real exports, real gross domestic product of importers and bilateral RMB real exchange rate, and the only times series of intra-regional RER flexibility between China and the rest of East Asia. First, we conduct Levin et al. (2002) panel unit root tests. The assumption of their model is that the cross-sectional time series are independently distributed and that the autoregressive parameter is identical for all cross-sections. The test allows for cross-section specific intercepts and/or time trend. Moreover, the error variance is also permitted to vary across the cross-sectional units. Their Monte Carlo simulation results showed that the tests had smallest size distortions and performed best against the homogenous alternative for panels of moderate size.

Panel A of Table 2 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 error processes. By contrast, for the intra-regional RER flexibility variable RER_{wt}, which is 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. The general finding is that the dependent variable (real exports) is trend stationary, and the set of regressors includes at least one trend stationary series, i.e., real GDP. There is no one series that contains both deterministic and stochastic trend components. These results suggest that we cannot model the variables $(\mathrm{y}_{it}, \mathbf{x}_{it})$ as a cointegrated system. The study also considers Pesaran (2007) cross-sectionally augmented Dickey-Fuller (CADF) regression that allows for panel heterogeneity and obtains CADF panel unit root test statistics. Panel B of Table 2 shows that y_{it} and GDP_{it} are trend stationary series with serially correlated errors, whereas the RMB bilateral real exchange rate and RER_{wt} are unit root processes. Again, since y_{it} and GDP_{it} are evidently trend stationary series, it is unlikely that the variables $(\mathrm{y}_{it}, \mathbf{x}_{it})$ are cointegrated. Standard estimation methods, such as dynamic OLS or fully-modified OLS, are thus quite likely to produce spurious results for the long-run parameters. So is the irrelevance of the empirical gravity model, which has been the workhorse to estimate the effect of border and/or currency union on trade integration. These specifications are restrictive because they presuppose a long-run equilibrium relationship in the observed data. The present study rather chooses a more general dynamic panel specification as outlined below.

4.2 The Dynamic Panel Data Model

In this study, China's export demand function is modeled in a dynamic framework, which is formulated as an autoregressive and distributed lag (ADL) model of order (2, 2)[15]:

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

[15] The selection of the order of autoregressive and distributed lag terms is based on Akaike's and Schwarz Bayesian Information Criteria.

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} the bilateral real exchange rate between China and country i which imports final exports from China (an increase denotes a real appreciation of the Chinese RMB), RER_{wt} the intra-regional RER flexibility between China and the rest of East Asia, 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[16]. 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 $\mathbf{d_{it}}$ indicates the deterministic variables (intercept and/or trend terms) and $\boldsymbol{\delta}_\mathbf{i}$ 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 export demand function in the imperfect substitutes framework (e.g., Chang, 1948; and Goldstein and Khan, 1985). According to the imperfect substitutes model, the observed demand function is the equilibrating behavior of both the supply-side and the demand-side of the model and, therefore, price-quantity relationship is, at least in theory, simultaneous. The empirical literature has taken the supply-side by assumption that the price elasticity of supply is infinite. This is restrictive given that exports production is increasingly fragmented across national borders in East Asia. Moreover, East Asian countries not only compete at each other's market, but more so in the Americas and Europe. This implies that demand schedules for exports from any East Asian country must be widely fluctuating for various factors. E. J. Working (1927) argued that if the demand curve did not shift much, but the supply curve did, then the intersection points would come to tracing a demand curve. He added that, by "correcting" for the influence of determinants, which cause demand curves to shift, one would obtain a better approximation of the true demand curve, even though the original demand schedules fluctuated widely. In the present context, it is rational to argue that RER flexibility among East Asian countries would have direct bearing on the demand schedules of China's exports. Therefore, our empirical specification (3) essentially includes the intra-regional RER flexibility variable, which is also the variable of interest, together with other covariates as suggested by the imperfect substitutes model.

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

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 (Hendry et al., 1984). 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).

4.3 Estimation Methods and Specification Tests

The empirical specification (3) is a dynamic error-component model. 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. In a dynamic panel model that includes unobserved fixed effects, the pooled OLS estimators are upward biased, because they are based on the restrictive assumptions that $\mathrm{E}(\mathbf{x}'_{\mathrm{it}}\mathrm{u}_{\mathrm{it}})=0$ and $\mathrm{E}(\mathbf{x}'_{\mathrm{it}}\eta_{\mathrm{i}})=0$, for $\mathrm{t}=1,\cdots,\mathrm{T}$. The dynamic model with lagged dependent variable must violate the assumptions because $\mathrm{y}_{\mathrm{it}-1}$ and η_{i} are correlated. Nickell (1981) showed that, for an autoregressive model that included a vector of truly exogenous variables, the within estimation of the autoregressive parameter would be downward biased, while the bias in the coefficient vector of the included exogenous variables would depend on the relationship between the exogenous variables and the lagged dependent variable y_{it-1}. Wooldridge (2002) showed that if u_{it} were correlated with future values of the explanatory variables in the sense that $\mathrm{E}(\mathbf{x}'_{\mathrm{it}}\mathrm{u}_{\mathrm{is}})\neq 0$ for $\mathrm{s}<\mathrm{t}$, the strict exogeneity assumption would fail in a dynamic panel model. And this will cause unknown bias in the fixed effect estimator. In addition, if the process $\{\mathbf{x}_{\mathrm{it}}\}$ has very persistent elements, the within estimator can also have substantial bias.

The present study, therefore, follows the Generalized Methods of Moments (GMM) approach for dynamic models of panel data as suggested by Holtz-Eakin et al. (1988), Arellano and Bond (1991), Arellano and Bover (1995), and Blundell and Bond (1998). In fact, the approach is a generalization of the IV estimation originally proposed by Anderson and Hsiao (1981 and 1982). For example, Anderson-Hsiao IV estimators of a dynamic panel model in first differences use either y_{it-2} or Δy_{it-2} as instruments for the lagged dependent variable Δy_{it-1}. By contrast, the GMM approach exploits further population moment conditions that can be related to both the differenced equations and the levels equations of the dynamic model.

Arellano and Bond (1991) considered a dynamic model $y_{it} = \alpha y_{it-1} + \boldsymbol{\beta}'\mathbf{x_{it}} + \eta_i + u_{it}$, where $\mathbf{x}_{it}$ is a $(K \times 1)$ vector of time-varying explanatory variables[17]. The basic assumption of their approach is that u_{it} have finite moments and, in particular, $E(u_{it}) = E(u_{it}u_{is}) = 0 \;\; \forall t \neq s$. That is, u_{it} are assumed to be serially uncorrelated. The model does not require any other knowledge concerning initial conditions or the distributions of the u_{it} and the η_i. In the first-differenced equations of the dynamic specification, the above assumptions lead to a set of linear moment conditions. However when $\mathbf{x}_{it}$ are assumed to be correlated with the unobserved fixed effects η_i, the optimal matrix of instruments crucially depends on whether the x_{it} are endogenous, predetermined or strictly exogenous. For example, if the x_{it} are endogenous in the sense that $E(\mathbf{x}'_{it}u_{is}) \neq 0$ for $s \leq t$ but zero otherwise, then x_{it} are treated symmetrically with the dependent variable y_{it}. In this case, the complete set of moment conditions available has the form of $E(Z'_i \Delta u_i) = 0$ for $i = 1, \cdots, N$, where $\Delta u_i = (\Delta u_{i3}, \cdots, \Delta u_{iT})'$ and the optimal matrix of the instruments $Z_i = \mathrm{diag}(y_{i1} \cdots y_{is} \mathbf{x}'_{i1} \cdots \mathbf{x}'_{is})$ $(s = 1, \cdots, T-2)$ are the valid instruments in the differenced equations. On the other hand, if the $\mathbf{x}_{it}$ are predetermined in the sense that $E(\mathbf{x}'_{it}u_{is}) \neq 0$ for $s < t$ but zero otherwise, the optimal matrix of the instruments $Z_i = \mathrm{diag}(y_{i1} \cdots y_{is} \mathbf{x}_{i1} \cdots \mathbf{x}_{is+1})$ are the valid instruments in the differenced equations. If we make much stronger assumption that the $\mathbf{x}_{it}$ are strictly exogenous, i.e., $E(\mathbf{x}'_{it}u_{is}) = 0 \;\; \forall s,t$, then the complete time series $\mathbf{x}'_i = (\mathbf{x}'_{i1}, \cdots, \mathbf{x}'_{iT})$ will be the valid instruments in each of the differenced equations. The optimal matrix is $Z_i = \mathrm{diag}(y_{i1} \cdots y_{is} \vdots \mathbf{x}'_i)$ for the period $(s = 1, \cdots, T-2)$.

Arellano and Bond (1991) thus suggested that lagged values of the dependent/endogenous variable itself and past, present and future values of the strictly exogenous variables would be valid instruments for the lagged dependent variable and other non-exogenous variables in the differenced equations of later period. GMM estimators that are based on moment conditions related to the differenced equations are referred to as the first-differenced GMM estimators. Let the expression $E(Z'_i \Delta u_i) = 0$ be the appropriate orthogonality conditions to be used to construct an estimator of the unknown parameter vector $\mathbf{B_0}$. Following Hansen (1982), the random function $g_N(\mathbf{B}) = N^{-1} \sum_{i=1}^{N} Z'_i \Delta u_i = N^{-1} \mathbf{Z}' \boldsymbol{\Delta} \mathbf{u}$ is the method of moments estimator of $E(Z'_i \Delta u_i)$, where A_N is a random weighting matrix . The GMM estimator $\hat{\mathbf{B}}$ is the set of elements in the parameter space that minimizes the sample criterion function $\left| h_N(\mathbf{B}) \right|^2$, where $h_N(\mathbf{B}) = A_N g_N(\mathbf{B})$ is the sample objective function . The first-order conditions of the minimization problem have the interpretation of setting k linear combinations of the

[17] The extension of the autoregressive specification to the case where a limited amount of serial correlation is allowed in u_{it} is straightforward.

r sample orthogonality conditions to zero where k is the dimensionality of the parameter space.

In a dynamic panel model r sample orthogonality conditions $g_N(\mathbf{B}) = 0$ often exceeds k parameters to be estimated. The weighting matrix $\mathbf{A_N}$ in fact reduces the number of equations to k by using linear combinations of r equations. Arellano and Bond (1991) showed that in the first-differenced equations of the dynamic panel model, the GMM estimator of the coefficient vector $\mathbf{B}' = (\boldsymbol{\alpha}', \boldsymbol{\beta}')$ is $\hat{\mathbf{B}}_{\text{diff}} = (\overline{\mathbf{X}}'\mathbf{Z}\mathbf{A_N}\mathbf{Z}'\overline{\mathbf{X}})^{-1}\overline{\mathbf{X}}'\mathbf{Z}\mathbf{A_N}\mathbf{Z}'\overline{\mathbf{y}}$, where $\overline{\mathbf{X}}$ is a stacked $(T-2)N \times K$ matrix of observations on $(\overline{y}_{it-1}, \overline{\mathbf{x}}_{it})$, $\overline{\mathbf{y}}$ and $\mathbf{Z}$ are accordingly defined for the appropriate choice of Z_i. The alternative choice of the weighting matrix $\mathbf{A_N}$ will give rise to GMM estimators with different asymptotic covariance matrices. For instance, one-step GMM estimators can be obtained by setting the weighting matrix $\mathbf{A}_N = (N^{-1}\sum_i Z_i' H Z_i)^{-1}$, where H is a $(T-2)$ square matrix with twos in the main diagonal, minus ones in the first sub-diagonals and zeros otherwise. On the other hand, one could obtain an "optimal" estimator of the weighting matrix from a family of random weighting matrices. An optimal weighting matrix is the one that has an asymptotic covariance matrix at least as small as any other element in the class. GMM estimator based on the optimal weighting matrix is called the two-step estimator. White (1982) suggested another choice of $\mathbf{A}_N$, which would be $\hat{\mathbf{V}}_N^{-1} = (N^{-1}\sum_i Z_i'\hat{\overline{v}}_i\hat{\overline{v}}_i' Z_i)^{-1}$, $\hat{\overline{v}}_i$ being the residuals from a preliminary consistent estimator of $\mathbf{B}' = (\boldsymbol{\alpha}', \boldsymbol{\beta}')$.

Arellano-Bond first-difference GMM estimators can, however, be further biased than the within estimators under certain conditions. Alonso-Borrego and Arellano (1996) and Blundell and Bond (1998) show that the first-differenced GMM estimators are weakly identified when the instruments are weak in the sense that they have a low correlation with the included endogenous variables. The estimators can be seriously downward biased in two important cases. First, as the value of the autoregressive parameter (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. In fact, when variables are persistent over time, lagged levels of these variables are weak instruments for the regression equation in differences. Weak instrument problem may not only cause the first-differenced GMM estimators to be further downward biased than the within estimators, it also influences the asymptotic and small-sample performance of the estimators.

To solve the problem, Arellano and Bover (1995) and Blundell and Bond (1998) proposed a new GMM estimator that would combine in a stacked system the regression in differences with the regression in levels. The instruments for the regression in differences are the same as suggested by Arellano and Bond (1991). But the instruments for the regression in levels are the lagged differences of the corresponding variables[18]. As

[18] It is based on an additional assumption that although there might be correlation between the levels of the right-hand side variables and the fixed effects, there would be no correlation between the differences of these variables and the fixed effects in dynamic panel model (see, e.g., Blundell Bond, 1998; and Levin et

the system of equations combines both the differenced equations and the levels equations, the instrument matrix is also an extended instrument matrix. The estimator based on this extended instrument matrix is called the system GMM estimator[19]. However, the choice between the differenced GMM estimator and the system GMM estimator is statistical and depends on whether there is strong persistency in the observed data.

We assess persistency characteristics of each individual time varying data series included in model (3). We estimate the univariate autoregressive model[20]

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

Here d_{it} is a vector of deterministic variables (e.g., intercept and/or time trend) and α_i is the corresponding vector of coefficients. For example, for the model with both intercepts and individual specific time trends, $d_{it} = \{1, t\}$. Here, α_{1i} represents 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 RER and RER_{wt} variables, we estimate the same autoregressive specification but without the trend element. 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 Table 2. The consistent GMM estimates of persistency are obtained by GMM system estimation.

Table 3 presents our results on persistency characteristics. 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 model 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 suggest that the multivariate dynamic panel data model in the present case must resolve on the potential weak instrument problem. The study therefore uses the extended instrument matrix as proposed by Blundell and Bond (1998) and obtains consistent the

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al., 2000). The assumption results from the stationary property that $E(y_{it+p}\eta_i) = E(y_{it+q}\eta_i)$ and $E(x_{it+p}\eta_i) = E(x_{it+q}\eta_i)$ for all p and q.

[19] Appendix 1 provides further details on the moment conditions and the resultant extended instrument matrices that are used to obtain GMM estimators under varying exogeneity assumptions.

[20] The form is the Sims et al. (1990) canonical form for higher order autoregressive processes, originally proposed by Fuller (1976).

system GMM estimators by estimating a system of differenced and levels equations of the dynamic specification.

As proposed in the conceptual framework, the paper estimates two benchmark specifications of model (3) in order to measure the impact on China's exports of a common currency in the region. The benchmark specifications are:

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

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

The specification (3.1) is 'the fully specified model,' while the specification (3.2) is 'the hypothetical model' based on the counterfactual assumption of a common currency in East Asia. The present study does not impose any arbitrary restrictions as to exogeneity of the included variables. In particular, we allow both RER_{cit} and RER_{wt} variables to be strictly exogenous, predetermined or endogenous and accordingly define the corresponding instrument matrices (see Appendix 1 for details). We thus obtain three sets of GMM estimators for both the specifications. In addition to the GMM-system estimators, we also report the pooled OLS and fixed-effect estimators in each case in order to show relative performance of the consistent system GMM estimators.

Several studies suggest that empirical export demand equation should include a supply shift variable. Hooper (1978) first argued that the observed high estimated income elasticity of demand for U.S. imports reflected the positive correlation between U.S. income growth and a relevant omitted variable, namely, supply capacity in the exporting countries, particularly the newly industrialized developing countries in East Asia. Since imports from these countries contained many new products with zero or unduly low weights in the standard price indices, the increased supply effect would not be reflected in recorded movements of U.S. import prices. Hooper thus suggested for including a supply proxy along with the normal arguments. In a similar vein, Chinn (2005) and 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. 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, since the country does not have consistent price index of exports. 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) found that

Hong Kong price index traced the price movement of China's exports better than others. It is highly plausible because of Hong Kong's traditional role as an entrepôt to transship China's exports to the rest of the world. Eichengreen et al. (2004) and Thorbecke (2006) applied the U.S. consumer price index to deflate U.S. dollar imports, 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 et al. (2006) used the U.S. Bureau of Labor Statistics (BLS) import price index of manufactured imports from non-industrial countries to deflate dollar value of China's exports. They found that the series closely matched the BLS price deflator for imports from China, which had been compiled since 2003. 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 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 Δu_{it} be the first differences of serially uncorrelated errors u_{it}. Then $E(\Delta u_{it} \Delta u_{i(t-1)})$ need not be zero, but the consistency of the GMM estimators fundamentally depends upon the assumption that $E(\Delta u_{it} \Delta u_{i(t-1)}) = 0$. 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 $(\mathrm{r}-\mathrm{k})$ 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[21]

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 Agency, Hong Kong.

Deflators: Hong Kong export price index is line 74d of IMF International Financial Statistics. Both the U.S. consumer price index and the import price index of manufactured imports from non-industrial countries are taken from the online database of the U.S. Bureau of Labor Statistics (BLS).

Real Exchange Rates: Bilateral real exchange rates between China and country i, i.e., RER_{ci}, and bilateral real exchange rates between China and country j that supplies intermediate goods to China, i.e., RER_{jc}, are all taken from the CHELEM database (CHEPII, 2007). In the calculation of $RER_{wt} = \Sigma\omega_{jt} RER_{jct}$, both RER_{wt} and ω_{jt} are annually updated over the sample period for each cross-section j. Note that RER_{wt} is a time series variable and hence uniform across cross-sections in the panel of China's bilateral 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, 2006; p. 7), and China's fixed capital formation (IFS series 93_E).

5. Results and Interpretation

Table 4 presents estimation results of the benchmark models for China's processing exports. These are the exports whose value chain is sharply fragmented across national borders, with final stages of assembly and exporting being performed in China. The first four columns report alternative estimation results, including the consistent system GMM estimates, of the fully specified model. Whereas, the fifth column reports only the system GMM estimates of the hypothetical model, which excludes the intra-regional RER flexibility as an explanatory variable. The parameters of interest are coefficients of both

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

the intra-regional RER flexibility and the bilateral RMB real exchange rate. 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 are obtained by using Hong Kong export price index as the deflator[22].

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. GMM1 assumes that both the RER variables are exogenous in the model, while GMM2 assumes that they are predetermined[23]. 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. Both GMM1 and GMM2 estimates 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 are largely comparable to the one-step estimates and, therefore, not reported. 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 tests of serial correlation in the first-differenced residuals are in both cases consistent with the maintained assumption of no serial correlation. However, the Sargan/Hansen test of over-identifying restrictions shows that the null of instrument validity is rejected for the GMM1 estimates, but not for the GMM2 estimates. Therefore, GMM2 specification of column 4 provides the consistent system GMM estimates of the fully specified model. Finally, column 5 of Table 4 reports GMM2 estimation results of the hypothetical model.

Here we focus on the contemporaneous effects because the study assumes that flexibility of exchange rates in East Asia is caused by asymmetry between exchange rates of national currencies, not by any real disequilibrium. The assumption does not preclude that deviations from long-run equilibrium would not have a long-run trade effect. The long-run estimates are discussed later. According to the preferred GMM2 estimates, the impact elasticity of $\mathrm{RER}_{\mathrm{w}}$ is −1.31 and that of $\mathrm{RER}_{\mathrm{ci}}$ is −0.75. The findings indicate that 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. By contrast, a 10 percent *flexibility* in relative prices between China and the other East Asian countries that supply intermediate goods to China will cause China’s processing exports to decline by

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[22] 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 by including alternative proxies to control for exporters’ increased capacity to supply new varieties.

[23] 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 endogeneity of a variable x_{it} requires that its lags dated $(t-2)$ or earlier and lagged differences dated $(t-1)$ are the valid instruments in the differenced equations and the levels equations respectively. If x_{it} is, in fact, not endogenous, the assumption is restrictive and will produce biased estimates.

13 percent. The term *flexibility* connotes misalignment in relative prices along the production networks in East Asia, rather than the traditional concept of exchange rate volatility. The finding shows that the impact of China's real exchange rate flexibility with the rest of East Asia on its processing exports is almost double the corresponding impact of a unilateral RMB appreciation.

Recall that our ultimate 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 stronger production network-linkage with the rest of East Asia. The above estimates are an intermediate step to that end.

To focus on the issue of not having a common currency, we refer to column 5 of Table 4. Column 5 reports estimates of the hypothetical model, which is based on the counterfactual assumption that there exists a common currency and that the RER_{w} variable is irrelevant as an explanatory variable. The finding is that the coefficient of RER_{ci} is now upward biased by the magnitude of 0.20 from the consistent estimate (i.e., −0.75) of the fully specified model. Now, we can identify two effects of the intra-regional RER flexibility on China's processing exports. One is its direct impact, i.e., $\beta_1 = -1.30$, while the other is its indirect impact on the coefficient of RER_{ci}. The indirect impact is estimated to be $(\beta_1 b_{12}) = -0.20$ [24]. 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 $(\beta_1 + \beta_1 b_{12}) = -1.50$ [25]. The first component is the RER_{w} coefficient, which measures the impact of real exchange rate misalignment between China and the rest of East Asia, whereas the second component is its indirect effect by increasing variability in the bilateral RMB real exchange rates.

Table 5 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 GMM2 to be the preferred estimators as shown in column 4. The reason is that there is no evidence of second-order serial correlation in the first-differenced residuals, that the null of instrument validity is not rejected and that the estimate of the autoregressive parameter is well within the range of the OLS and FE estimates. The results show that a 10 percent unilateral RMB appreciation against the rest of the world currencies will cause China's ordinary exports to decline by 8.9 percent. On the other hand, a 10 percent *flexibility* in relative prices between China and the other East Asian countries will cause the ordinary exports to decline by 6.5 percent. This indicates that a unilateral RMB appreciation would have larger impact on the ordinary exports than the processing exports. A relatively larger coefficient of the RMB RER is consistent with the fact that the extent of local value addition is substantial in the gross value of China's

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[24] The notations β_1, β_2 and b_{12} used in this section are consistent with the notations introduced in the conceptual framework of the paper, not the coefficients of the estimated model.

[25] We estimate standard error of $\beta_1 + \beta_1 b_{12}$ by using delta method and find they are significant at 1 % level.

ordinary exports. By contrast, the impact of intra-regional RER flexibility on the ordinary exports is just half of the corresponding impact on the processing exports. A relatively lower magnitude of RER_{w} coefficient signifies the weak linkage of the ordinary exports with East Asian production networks.

By the same analogy, we find that the cost of not having a common currency area in East Asia is modest in the case of China's ordinary exports. Column 5 of Table 5 shows that the elasticity of RER_{ci} in the hypothetical model is upward biased by a magnitude of 0.10. Therefore, the combined cost of not having a common currency for the ordinary exports is $(\beta_1 + \beta_1 b_{12}) = -0.75$. This is just half of the cost we have estimated in the case of China's processing exports.

The findings carry important implications for East Asian production networks and China's exports. Earlier we observed that *'processing trade'* is at the heart of China's integration with the rest of East Asia and to the world trading system. The results in Table 4 indicate that flexibility in East Asian exchange rates, due to either an appreciation or a depreciation of world invoice currencies, does greatly affect China's processing exports. It does it by misaligning the relative price relationships between China and the rest of East Asia, not between East Asia and the rest of the world. This is what the coefficient of RER_w variable implies. The finding shows that the impact of RER_{w} on the processing exports is twice as much as that of a unilateral RMB appreciation. The reason is that a unilateral RMB appreciation affects only the Chinese value added, whereas a relative price misalignment between China and the rest of East Asia affects the dollar costs of intermediate goods imported into China from the rest of East Asia[26]. Independent national currencies linked by flexible exchange rates not only cause exchange rate uncertainty but also impede myriad contractual arrangements related to international trade. The resultant effect of rising trade costs would significantly affect the processing exports, not the ordinary exports. A comparison of the opportunity costs of not having a regional currency in terms of processing-and-ordinary decomposition is reflective of this point. The cost to China's processing exports is just double the corresponding cost to China's ordinary exports. The results also imply that contemporaneous effect of a discrete exchange rate shock can be prohibitively high for the processing exports, and since they are produced regionally, the effect would be contiguous in nature.

We now turn to the estimates of long-run parameters and their test statistics that are based on preferred GMM2 estimates of the fully specified model (3.1). With respect to RER_{ci} and RER_{w}, the respective long-run elasticities are $\hat{\xi} = (\hat{\xi}_0 + \hat{\xi}_1 + \hat{\xi}_2)/(1 - \hat{\alpha}_1 - \hat{\alpha}_2)$ $\hat{\psi} = (\hat{\psi}_0 + \hat{\psi}_1 + \hat{\psi}_2)/(1 - \hat{\alpha}_1 - \hat{\alpha}_2)$. For the panel of processing exports, Table 6 shows that the long-run effect of a unilateral RMB appreciation is merely −0.93 and that of the

[26] Lau and Stiglitz (2005) argued that the dollar costs of imported intermediate goods would be more than two-thirds of the gross value of China's processing exports.

intra-regional RER flexibility variable is −8.5[27]. In the long-run, we do not find any significant evidence of indirect effect as earlier measured by the term $(\beta_1 b_{12})$ in the analysis of short-run dynamics. The study also finds that the long-run parameter estimates for the panel of ordinary exports are statistically insignificant, when they are based on the consistent dynamic estimates obtained by using GMM system estimators. The implication is that flexible exchange rates between national currencies and fragmentation of production process across borders in East Asia are incompatible in the long-run.

Finally, the study conducts comparative static implications of intra-regional RER flexibility, which is otherwise the lack of a common currency arrangement in East Asia, for China's bilateral processing exports. It is because the long-run effect of RER_w is statistically significant only for the processing exports. Table 7 reports the findings. Column 3 shows mean differences between potential and actual volume of processing exports for each bilateral trading partner. The averages are taken over the 1994-2005 period[28]. Column 4 shows t-ratios testing if these mean differences are significantly greater than zero. The last column shows ratios between actual and potential volume of exports at their respective averages.

Major finding is that the production and exporting of processing goods is, on an average over the sample period, 20 percent below the potential. At a disaggregated importer level, the extent of average trade loss is the highest for Japan ($15 billion), followed by the US ($5.5 billion), European Union ($5 billion), Taiwan ($2.5 billion) and South Korea ($2.2 billion). The trade loss in the case of Hong Kong, being an entrepôt to the west, should better be regarded as trade loss against the Americas and Europe. The finding that the actual exports volume is far below the potential level for Japan and NIEs probably indicates that the intra-regional exchange rate asymmetry has constrained the development of East Asian production networks to reach its optimal degree. Given a fixed exchange rates system, the production networks would further evolve to fragment production processes for a larger class of goods. It would lead to an increasing level of back-and-forth trade in intermediate goods between countries but along the production networks. The networking would be both vertical and horizontal in order to take advantage of differences in technologies, factor endowments and market sizes across countries.

The deepening integration of East Asian countries with one another, which is mostly driven by market forces, has produced new demand for a regional currency and against national currencies. However, East Asian countries have their independent national

[27] Standard error for the estimator of the long-run parameter is obtained by using delta method. Pesaran and Shin (1998) showed that variance estimator obtained by the delta method is asymptotically valid irrespective of whether x_t is I(1) or I(0). The test statistics indicate that the estimates of the long-run parameters are statistically significant at 5%.

[28] To estimate average trade effect over T is somewhat less meaningful for policy analysis because any real misalignment would be discrete, should there be an external shock. An average understates the effect of a crisis period. Though it is done for analytical ease, the estimates would probably capture the effect of persistent real misalignment due to RER heterogeneity within East Asia.

currencies and pursue heterogeneous exchange rate and monetary policies. The evidence suggests that the opportunity cost of not having a regional currency can be prohibitive for those exports that are produced along the regional production networks. The growing pattern of East Asian integration is, therefore, very susceptible to external shocks, e.g., a discrete depreciation of the world invoice currencies like the U.S. dollar or the Euro.

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 to the inclusion of additional control variables in the estimated fully specified model[29].

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. Columns 1 and 2 of Table 8 show the key parameter estimates for the case of China's processing exports. The coefficient of the RMB real exchange rate is −0.75, while that of the intra-regional RER flexibility is in the range of −1.08 to −1.33. Columns 3 and 4 show the corresponding results for China's ordinary exports. The results show that the coefficient of RER_{ci} is about −0.90, while that of RER_{wt} is very low in magnitudes and statistically insignificant when the deflator is the U.S. CPI. Otherwise, all these estimates are statistically significant at any reasonable level. The estimates are within one standard error of the corresponding estimates for the benchmark results that are reported in Tables 4 and 5 respectively. In other words, the estimates of the impact of real exchange rate misalignment and of the costs of not having a common currency on China's exports, particularly the processing exports, do remain robust, regardless of how we deflate nominal dollar value of the exports. The long-run behavior of the model is also unchanged form the benchmark case.

Next, we augment the benchmark models by including a supply shift variable to control for structural break in the estimated relationships and to examine if the estimates of the key parameters are robust to the inclusion of the variable. Since it is difficult to find a good proxy for the supply shift effect, we follow the recent trade literature and use alternative proxies such as real GDP of China, cumulative foreign direct investment (FDI) into China and China's gross fixed capital formation. It should be noted that all

[29] All the estimates that are reported in this section are one-step system GMM estimators that are obtained under the assumption that RER_{ci} and RER_{w} are predetermined (weakly exogenous) in the model. Though the pooled OLS, fixed-effect and differenced GMM estimators are relatively biased for the reasons explained in the methodologies, the key parameter estimates under these alternative estimators are nonetheless comparable with those of the consistent system GMM estimates.

proxies are *too* endogenous. The study thus uses the lag dated $(t-2)$ of the proxy as an argument in the model. The study does not use China's capital stock as a proxy, because the variable is both endogenous and measured with much error. Instead, the gross fixed capital formation will be a better proxy, since it is correlated with the capital stock but uncorrelated with the measurement error.

Columns 5 through 10 of Table 8 report estimates of the key parameters of the augmented model. The results show that the inclusion of proxy for the supply side effect significantly affect the short-run dynamics of the intra-regional RER flexibility between China and the rest of East Asia. But the direction of the effect is invariably downward to be further negative. This is true for both the processing exports and the ordinary exports, though their magnitudes are different as expected. The findings clearly imply that the benchmark estimates that are reported respectively in Tables 4 and 5 (column 4 in each case) provide a conservative lower-bound limit for the estimates of the key parameters. The long-run behavior of the model is also unchanged that the long-run effect of intra-regional RER flexibility would be prohibitively high for exports that are produced along the production networks.

In the view of a small sample involving 34 cross-sections and 14 periods, we consistently estimated the autoregressive and distributed lag model of the second order. As we indicated earlier, the selection of the order of autoregressive and distributed lags is based on the minimization of Akaike's and Schwarz Bayesian Information Criteria. An arbitrary first order ADL specification with (or without) the proxy for the supply shift effect causes no change in the short-run dynamics, but only causes the long-run effect of RER_w to be −2.87. The key parameter estimates are statistically significant at 1%. Furthermore, 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 also provide general support for the ADL(2,2) model used in the present study.

7. Concluding Remarks and Directions for Future Research

In this paper, we have developed a framework for assessing the impact of a common currency on East Asian production networks and China's exports behavior. The framework produces a new variable in order to capture the asymmetry in exchange rate and monetary policies between China and other East Asian countries that supply intermediate goods to China. The variable is constructed by using the degree of production network linkage of China's final exports as the weight on the level of real exchange rate misalignment between China and another East Asian country. The overall framework has the utility to be applied to estimate *ex ante* effect of a common currency on trade integration.

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 long-run effect of the intra-regional RER flexibility on the processing exports is almost 9.0 times the corresponding estimate of a unilateral appreciation of China's RMB exchange rate. The magnitudes of these estimates 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 increasingly fragmented across borders in East Asia.

The major limitation is that the study focuses on a very narrow domain. It investigates only trade effect of a common currency proposition on East Asian production networks, particularly focusing on China's processing exports. To the extent export production of other East Asian countries is too integrated with the production networks, it is expected that a fixed exchange rate system would raise their export potential in general. However, it is imperative to investigate spatial distribution of welfare implications of the policy. Secondly, in the literature there has been a disconnect between trade effect of exchange rate uncertainty (as measured by shot-run volatility) and that of a common currency arrangement. Though the present study offers a viable solution of this disconnect, it does not offer a coherent theoretical framework. Future research can be directed to examine how exchange rate flexibility enters into firm's decision-making process, when production of a value chain is organized across national borders. Thirdly, the econometric modeling applied in the present study is though unconstrained and not forced to obey some economic theory, it requires further refinements for drawing valid inference on theory-consistent long-run relations. Since the observed data characterize temporal aggregation across countries, inference on the long-run relations based on a finite-order ADL model may not be valid. Further investigation is necessary as to panel unit roots and cointegration relations. However, it is hoped that both the conceptual framework and estimation methodologies will motivate further empirical research on trade effect of a common currency arrangement in East Asia.

Appendix-1

Moment Conditions and the Instrument Matrix

This appendix provides details on the moment conditions and related GMM-style instrument matrices for obtaining system GMM estimators. For convenience, we omit the time-invarying set of variables, though they are relevant instruments in the levels equations and have been used in the system GMM estimation. Let us recast our ADL (2,2) model (3) without the time-invarying variables:

$$y_{it} = \sum_{k=1}^{2} \alpha_k y_{it\text{-}k} + \sum_{k=0}^{2} \boldsymbol{\beta}_{\mathbf{k}}' \mathbf{x}_{\mathbf{it-k}} + \eta_i + \boldsymbol{\delta}_{\mathbf{i}}' \mathbf{d}_{\mathbf{it}} + u_{it}, \quad t = 3,\cdots,T;\ i = 1,\cdots,N. \qquad (3^*)$$

Below we define how both the moment conditions and the resultant extended instrument matrix depend on varying exogeneity assumptions regarding the vector $\mathbf{x}_{\mathbf{it}}$, provided that y_{t-1} is always endogenous in the first differenced equations.

1. GMM1 Estimator [Assumption: The vector $\mathbf{x}_{it}$ is strictly exogenous in the model, i.e., $E(\mathbf{x}_{\mathbf{it}} u_{is}) = 0\ \forall s, t$]

i. Moment conditions for the differenced equations: $\mathrm{E}(\mathrm{y}_{\mathrm{it-s}} \Delta \mathrm{u}_{\mathrm{it}}) = 0$ and $E(\mathbf{x}_{\mathbf{i}} \Delta u_{it}) = 0$ where $\mathbf{x}_{\mathbf{i}}' = (\mathbf{x}_{\mathbf{i1}}' \cdots \mathbf{x}_{\mathbf{iT}}')$ for $\mathrm{t} = 4,\cdots,\mathrm{T}$ and $\mathrm{s} \geq 2$. This implies that lags of y dated $(t-2)$ and earlier and past, present and future values of the exogenous variables are valid instruments for the lagged dependent variable in the differenced equations for $\mathrm{t} = 4,\cdots,\mathrm{T}$. The resultant optimal instrument matrix is $Z_i = diag(y_{i1}\ y_{i2} \cdots y_{is} \mathbf{x}_{\mathbf{i1}} \cdots \mathbf{x}_{\mathbf{iT}})$ for $\mathrm{t} = 4,\cdots,\mathrm{T}$ and $\mathrm{s} \geq 2$. Arellano and Bond (1991), however, did not use all the over-identifying restrictions arising from the strict exogeneity assumption of $\mathbf{x}_{it}$, but only the present values in their Monte Carlo experiments. Thus their instrument matrix looked like $Z_i = diag(y_{i1}\ y_{i2} \cdots y_{is} \vdots (\mathbf{x}_{\mathbf{i4}}' \cdots \mathbf{x}_{\mathbf{iT}}')')$ for $\mathrm{t} = 4,\cdots,\mathrm{T}$ and $\mathrm{s} \geq 2$. This instrument matrix is usually used to obtain the differenced GMM estimator, which will be inconsistent and inefficient in the presence the weak instrument problem. Arellano and Bover (1995) and Blundell and Bond (1998) therefore suggested a further set of moment conditions for equations in levels.

ii. The moment conditions for the levels equations: $\mathrm{E}(\Delta \mathrm{y}_{\mathrm{it-1}} \mathrm{u}_{\mathrm{it}}) = 0$ for $\mathrm{t} = 4,\cdots,\mathrm{T}$. Given that lagged levels are used as instruments in the difference equations, only most recent lagged difference is the valid instrument in the levels equations. Using the other lagged differences would results in redundant moment conditions.

Calculation of the system GMM estimators is essentially based on a stacked system comprising both the differenced equations and the levels equations of the model for $\mathrm{t} = 4,\cdots,\mathrm{T}$. The instrument matrix for this system is called the extended instrument matrix, which can be written as,

$$Z_i^+ = \begin{bmatrix} Z_i & 0 & 0 & \cdots & 0 \\ 0 & \Delta y_{i3} & 0 & \cdots & 0 \\ 0 & 0 & \Delta y_{i4} & \cdots & 0 \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & 0 & \cdots & \Delta y_{i,T-1} \end{bmatrix}, \text{ where } Z_i = diag(y_{i1}\ y_{i2} \cdots y_{is} \vdots (\mathbf{x}'_{\mathbf{i4}} \cdots \mathbf{x}'_{\mathbf{iT}})'.$$

2. *GMM2 Estimator [Assumption: The vector* $\mathbf{x}_{it}$ *is predetermined in the model, i.e.,* $E(\mathbf{x}_{\mathbf{it}} u_{is}) \neq 0$ for $s < t$ but zero otherwise.*]*
The moment conditions for the differenced equations are $\mathrm{E}(\mathrm{y}_{\mathrm{it-s}} \Delta \mathrm{u}_{\mathrm{it}}) = 0$ and $E(\mathbf{x}_{\mathbf{it-s+1}} \Delta u_{it}) = 0$ for $\mathrm{t} = 4, \cdots, \mathrm{T}$ and $\mathrm{s} \geq 2$. Whereas, for the levels equations, the additional moment conditions are $\mathrm{E}(\Delta \mathrm{y}_{\mathrm{it-1}} \mathrm{u}_{\mathrm{it}}) = 0$ and $E(\Delta \mathbf{x}_{\mathbf{it}} u_{it}) = 0$ for $\mathrm{t} = 4, \cdots, \mathrm{T}$. The only difference from the case 1 will be in the context of underlying extended GMM instrument matrix. While Z_i is $[diag(y_{i1}\ y_{i2} \cdots y_{is} \mathbf{x}'_{\mathbf{i1}} \cdots \mathbf{x}'_{\mathbf{is+1}})]$, Z_i^+ will incorporate additional instruments $[diag(\Delta y_{iT-1} \mathbf{\Delta x}'_{\mathbf{iT}})]$ for levels equations, where $\mathrm{t} = 4, \cdots, \mathrm{T}$ and $\mathrm{s} \geq 2$.

3. *GMM3 Estimator [Assumption: The vector* $\mathbf{x}_{it}$ *is endogenous in the model, i.e.,* $E(\mathbf{x}_{\mathbf{it}} u_{is}) \neq 0$ for $s \leq t$ but zero otherwise.*]*
In this case, the moment conditions for the differenced equations are $E(y_{it-s} \Delta u_{it}) = 0$ and $E(\mathbf{x}_{\mathbf{it-s}} \Delta u_{it}) = 0$ for $\mathrm{t} = 4, \cdots, \mathrm{T}$ and $\mathrm{s} \geq 2$. Whereas, for the levels equations, the additional moment conditions are $\mathrm{E}(\Delta \mathrm{y}_{\mathrm{it-1}} \mathrm{u}_{\mathrm{it}}) = 0$ and $E(\Delta \mathbf{x}_{\mathbf{it-1}} u_{it}) = 0$ for $\mathrm{t} = 4, \cdots, \mathrm{T}$. Again, calculation of the system GMM estimators is essentially based on a stacked system comprising both the differenced equations and the levels equations of the model. The extended instrument matrix Z_i^+ will now represent $Z_i = diag(y_{i1}\ y_{i2} \cdots y_{is} \mathbf{x}'_{\mathbf{i1}} \cdots \mathbf{x}'_{\mathbf{is}})$ being appended by the instrument set $[diag(\Delta y_{iT-1} \mathbf{\Delta x}'_{\mathbf{iT-1}})]$ for $\mathrm{t} = 4, \cdots, \mathrm{T}$ and $\mathrm{s} \geq 2$.

Note that potential endogeneity of a variable x_{it} requires that its lags dated $(\mathrm{t}-2)$ or earlier can only be the instruments for the differenced equations, while its lagged differences dated $(t-1)$ will be the valid instruments for the levels equations. If x_{it} is, in fact, not endogenous, the assumption is restrictive and causes biasness in the estimates. The GMM3 estimates, which 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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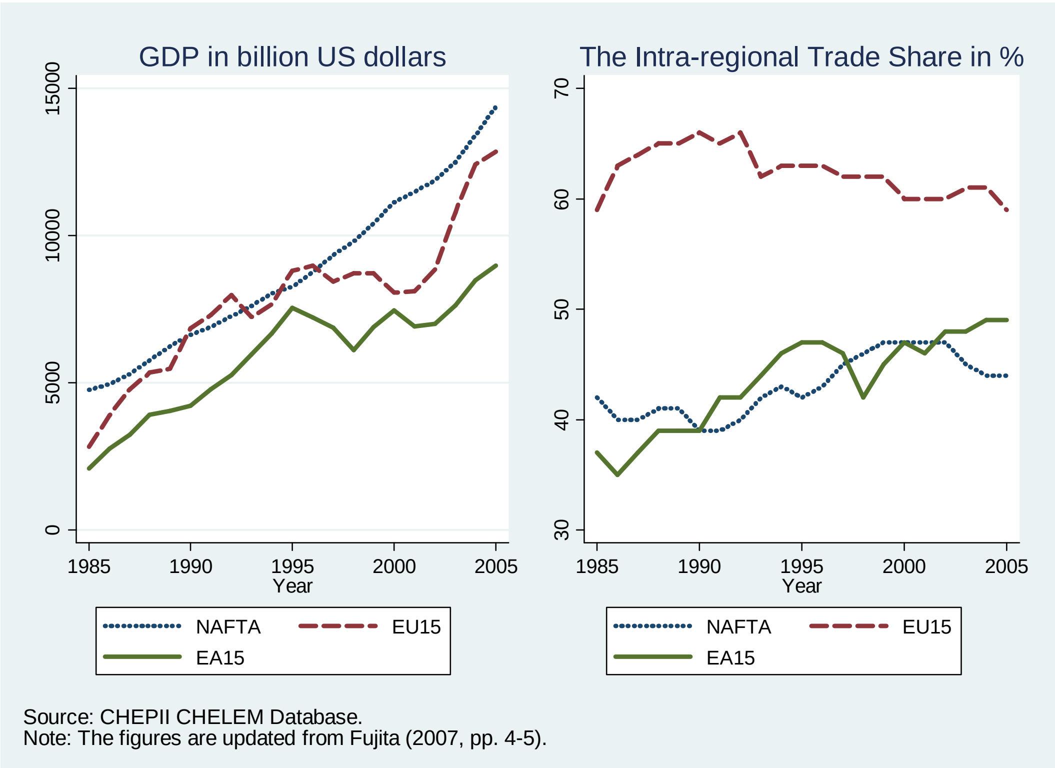

Figure 1. Globalization, regional growth and economic interdependency, 1985-2005

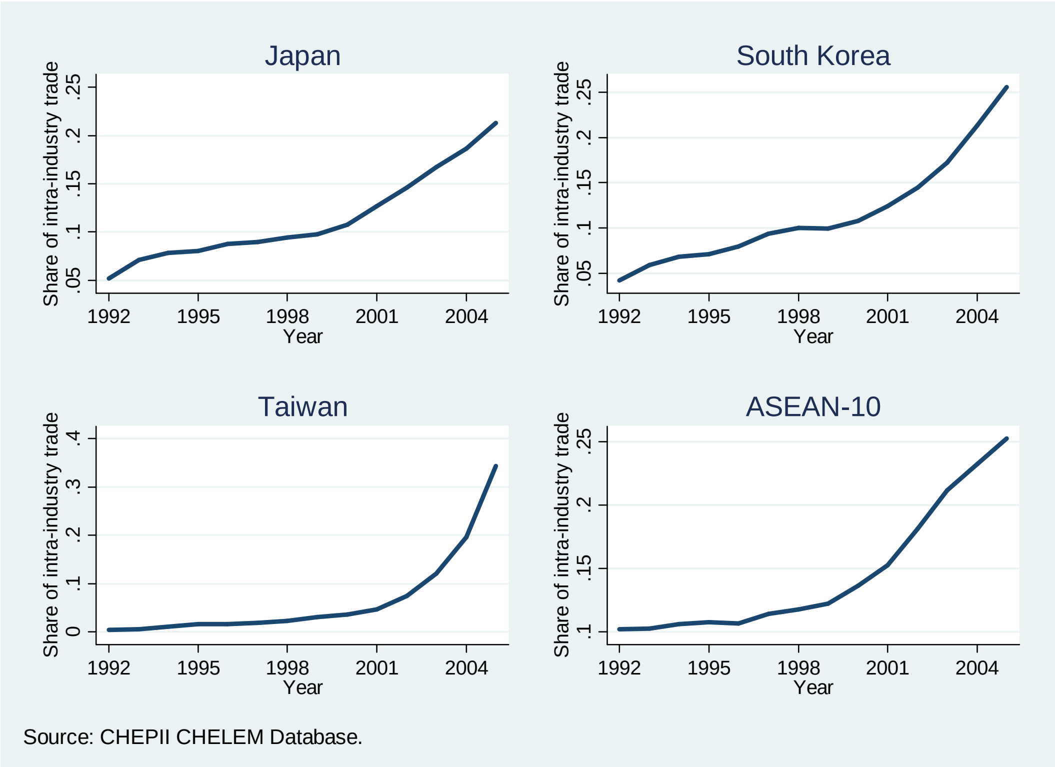

Figure 2. Intra-industry trade intensity between China and the rest of East Asia

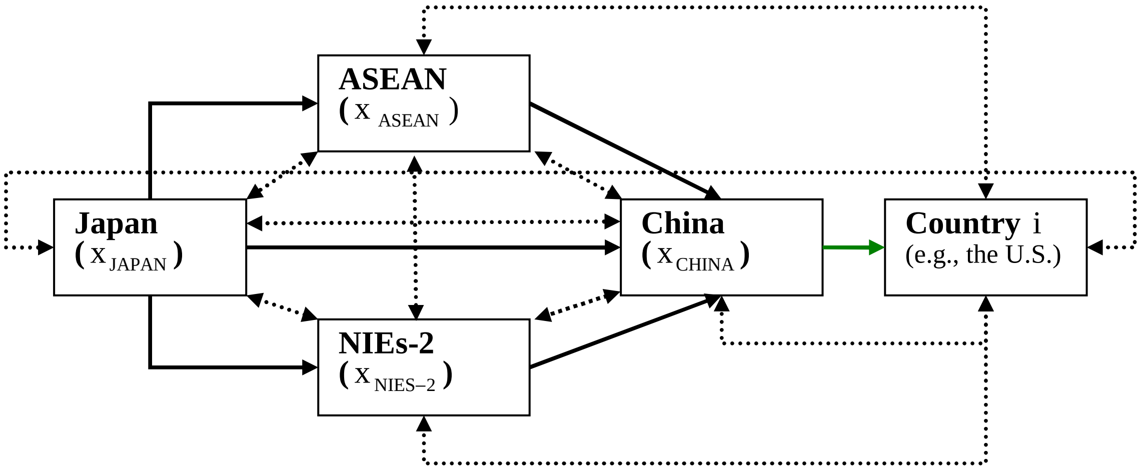

Figure 3. Schematic view of production networks and real exchange rate relationships

Table-1

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
a. Ordinary imports	36.6	7.9	2.1	3.3	1.1	14.3	5.1	7.9	9.3
b. Processing imports	35.0	7.7	10.8	1.9	7.0	27.3	1.9	1.7	4.0
c. 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
a. Ordinary imports	42.4	5.4	5.7	3.1	0.5	14.8	3.9	6.4	17.3
b. Processing imports	41.5	6.9	14.4	5.7	1.2	28.1	1.9	1.8	9.7
c. 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
a. Ordinary exports	47.1	9.8	2.7	3.6	16.0	9.6	5.8	6.8	8.9
b. Processing exports	48.2	7.3	2.0	1.4	10.7	14.0	12.7	6.5	4.3
c. 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
a. Ordinary exports	41.3	4.4	3.2	2.9	10.5	3.3	6.9	7.4	13.2
b. Processing exports	54.7	6.5	3.4	3.2	13.1	12.2	13.9	9.5	6.0
c. Other processing exports	4.0	0.1	0.2	0.2	0.5	0.9	0.6	0.4	1.7

Table-1 (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
a. Ordinary trade	5.2	0.7	0.3	-0.1	0.9	7.7	0.0	-2.0	5.7	-1.5
b. Processing trade	7.9	-1.3	-9.4	-0.6	-11.4	5.7	9.7	4.2	19.5	-0.3
c. 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
a. Ordinary trade	35.4	-2.5	-12.9	2.0	**-13.4**	21.6	26.9	14.4	**62.9**	-14.0
b. Processing trade	142.5	4.5	-69.3	-13.3	**-78.1**	85.1	92.9	60.4	**238.4**	-17.9
c. Other processing trade	-75.9	-18.5	-17.7	-12.4	**-48.6**	5.6	-5.6	-13.4	**-13.4**	-13.9

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. However, China's exports via Hong Kong, largely finished exports, are generally destined for the U.S. and EU-15 markets. Following Kwan (2002), China's bilateral trade surplus against Hong Kong is therefore considered as China's bilateral trade surplus against the U.S. and EU-15. 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. Trade balance on account of the "processing trade" is thus related to foreign affiliates of multinationals, while trade balance on account of both the "ordinary trade" and the "other processing trade" is related to Chinese-owned local firms. 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 and Thailand.

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

Table 2
Panel unit root tests

Panel A: Levin-Lin-Chu pooled augmented Dickey-Fuller (ADF) tests		
Test Statistic Variables	Levin-Lin-Chu ADF t_δ^* test	Specifications for deterministics and autoregressive order
Real exports (ordinary)	-11.336***	Constant and trend; AR(1)
Real exports (processing)	-16.065***	Constant and trend; AR(1)
Real GDP (GDP_{it})	-6.897***	Constant and trend; AR(1)
RMB RER (RER_{cit})	-2.749***	Constant; AR(4)
Panel B: Pesaran cross-sectionally augmented Dickey-Fuller (CADF) tests		
Test Statistic Variables	CADF $\bar{t}$ test	Specifications for deterministics and autoregressive order
Real exports (ordinary)	-2.522**	Constant and trend; AR(2)
Real exports (processing)	-2.599**	Constant and trend; AR(2)
Real GDP (GDP_{it})	-4.231***	Constant and trend; AR(2)
RMB RER (RER_{cit})	-0.679	Constant; AR(2)

A three-step procedure is followed to obtain Levin-Lin-Chu panel unit root tests. First, Δy_{it} and y_{it-1} are regressed on Δy_{it-L} $(L = 1, \cdots, p_i)$ for generating orthogonalized residuals $\hat{e}_{it}$ and $\hat{v}_{it-1}$ respectively. Second, the ratio of long run to short run innovation standard deviations for each cross-sectional unit is estimated. Finally, 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 estimate of the average standard deviation ratio is then used to adjust t_δ statistic from the above estimation to derive adjusted t_δ^* statistics. By contrast, Pesaran CADF $\bar{t}$ statistic is defined as $\bar{t} = N^{-1}\Sigma t_i$, where t_i is the cross-sectionally augmented Dickey-Fuller statistic for the i^{th} cross-section unit given by the t-ratio of the coefficient of y_{it-1} in the CADF regression. RER_{cit} represents the bilateral real exchange rate of Chinese renminbi vis-à-vis country i. The intra-regional RER flexibility (RER_{wt}) between China and other East Asian countries that supply intermediate goods to China 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. '*', '**' and '***' denote 10%, 5% and 1% statistical significance, respectively.

Table 3
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*** (0.013)	0.928*** (0.043)	0.719*** (0.063)	0.811*** (0.091)
Processing real exports ($EX2_{it}$)	0.979*** (0.009)	0.912*** (0.027)	0.754*** (0.060)	0.360*** (0.097)
Real GDP (GDP_{it})	0.997*** (0.001)	0.974*** (0.008)	0.771*** (0.043)	0.521*** (0.157)
RMB RER (RER_{cit})	0.999*** (0.018)	0.786*** (0.081)	0.282*** (0.094)	-0.070 (0.132)
Weighted RER (RER_{wt})	.881*** (.091)	n/a	n/a	n/a

The data series $EX1_{it}$ and $EX2_{it}$ represent China's bilateral ordinary and processing exports respectively to country i, GDP_{it} real gross domestic product of importing country i, RER_{cit} the bilateral real exchange rate between China and country i, RER_{wt} the intra-regional RER flexibility between China and countries that supply intermediate goods to to China. The OLS and within estimates of the parameter are biased upwards and downwards respectively. GMM first-difference estimates are further biased than the within estimates. The consistent system GMM estimator is obtained by estimating a system combining both the differenced equations and the levels of the univariate dynamic panel model. The moment conditions are $E(y_{it-s}\Delta e_{it}) = 0$ for $t = p+1,\cdots,T$ and $s \geq 2$, for the differenced equations, and $E(e_{it}\Delta y_{it-1}) = 0$ for $t = p+1,\cdots,T$, for the levels equations of the model. The above moment conditions give rise to the extended instrument matrix used to obtain the system GMM estimator. Only the OLS estimate of persistency of RER_{wt} is reported, since the variable does not vary cross-sectionally. '*', '**' and '***' denote 10%, 5% and 1% statistical significance, respectively.

Table 4
Estimation of autoregressive and distributed lag model for China's processing exports to 33 countries, 1992-2005 (Benchmark results; dependent variable: China's bilateral processing exports)

	Fully specified model				Hypothetical model
	1	2	3	4	5
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2	GMM2
Lagged real exports$_{i(t-1)}$	0.988***	0.774***	0.791***	0.776***	0.663***
	(0.073)	(0.084)	(0.069)	(0.073)	(0.090)
GDP of importer$_{it}$	2.466***	2.480***	2.674***	2.560***	2.494***
	(0.487)	(0.484)	(0.556)	(0.508)	(0.574)
GDP of importer$_{i(t-1)}$	-3.037***	-2.382***	-2.515***	-2.091**	-1.465**
	(0.770)	(0.739)	(0.715)	(0.797)	(0.704)
GDP of importer$_{i(t-2)}$	0.591	0.646	0.014	-0.297	-0.844*
	(0.452)	(0.451)	(0.392)	(0.521)	(0.482)
Bilateral RMB RER$_{ci,t}$	-0.784***	-0.799***	-0.718***	-0.754***	-0.584***
	(0.191)	(0.187)	(0.186)	(0.182)	(0.160)
Bilateral RMB RER$_{ci,(t-1)}$	0.523**	0.380	0.419*	0.429	0.178
	(0.252)	(0.231)	(0.246)	(0.263)	(0.167)
Bilateral RMB RER$_{ci,(t-2)}$	0.267*	0.082	0.254**	0.160	0.236**
	(0.157)	(0.155)	(0.115)	(0.165)	(0.088)
Intra-regional RER flexibility$_{wt}$	-1.691***	-1.368***	-1.088***	-1.306***	
	(0.436)	(0.270)	(0.224)	(0.368)	
Intra-regional RER flexibility$_{w(t-1)}$	0.770**	0.763**	0.598*	0.533	
	(0.351)	(0.340)	(0.295)	(0.340)	
Intra-regional RER flexibility$_{w(t-2)}$	-0.799**	-0.738***	-0.439*	-0.715*	
	(0.324)	(0.249)	(0.217)	(0.352)	
m1			-2.94***	-2.91***	-2.81***
m2			-0.29	-0.28	-1.11
Hansen J Statistic			0.009	0.590	0.111
P-value (d.f.)			(10)	(28)	(21)
No. of Groups	33	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396	396

Hong Kong export price index is used to deflate the nominal dollar value of the processing exports. Columns 1-4 report estimation results of the fully specified model, while column 5 reports estimation results of the hypothetical model, which excludes $RER_{w,(t-j)}$ by assumption. The coefficient of lagged real exports$_{i(t-2)}$ is always statistically insignificant. Both GMM1 and GMM2 are one-step system GMM estimates that are obtained by estimating a system of the differenced equations and the levels equations of the model. While GMM1 assumes that RER_{ci} and RER_w are exogenous, GMM2 assumes that they are predetermined in the model. 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. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses. Significance tests: *** p<0.01, ** p<0.05, * p<0.1.

Table 5

Estimation of autoregressive and distributed lag model for China's ordinary exports to 33 countries, 1992-2005 (Benchmark results; dependent variable: China's bilateral ordinary exports)

	Fully specified model				Hypothetical model
	1	2	3	4	5
Independent Variables	Pooled OLS	Fixed-Effect	GMM1	GMM2	GMM2
Lagged real exports$_{i(t-1)}$	0.657***	0.504***	0.572***	0.537***	0.545***
	(0.104)	(0.096)	(0.078)	(0.105)	(0.080)
Lagged real exports$_{i(t-2)}$	0.288***	0.198**	0.342***	0.384***	0.301**
	(0.096)	(0.087)	(0.089)	(0.103)	(0.111)
GDP of importer$_{it}$	1.795***	1.823***	1.768***	1.564**	1.850***
	(0.466)	(0.487)	(0.628)	(0.635)	(0.598)
GDP of importer$_{i(t-1)}$	-2.185***	-1.796***	-1.703**	-1.369	-1.820**
	(0.718)	(0.674)	(0.793)	(0.943)	(0.741)
GDP of importer$_{i(t-2)}$	0.449	0.039	0.026	-0.120	0.130
	(0.422)	(0.404)	(0.289)	(0.358)	(0.315)
Bilateral RMB RER$_{ci,t}$	-0.861***	-1.001***	-0.821***	-0.889***	-0.749***
	(0.128)	(0.139)	(0.160)	(0.150)	(0.118)
Bilateral RMB RER$_{ci,(t-1)}$	0.526***	0.417***	0.500**	0.480***	0.204
	(0.154)	(0.147)	(0.188)	(0.169)	(0.142)
Bilateral RMB RER$_{ci,(t-2)}$	0.414***	0.233*	0.411***	0.352**	0.559***
	(0.119)	(0.140)	(0.109)	(0.131)	(0.136)
Intra-regional RER flexibility$_{wt}$	-0.853***	-1.028***	-0.455**	-0.676**	
	(0.315)	(0.246)	(0.214)	(0.328)	
Intra-regional RER flexibility$_{w(t-1)}$	0.409*	0.348	0.436*	0.398*	
	(0.225)	(0.231)	(0.224)	(0.225)	
Intra-regional RER flexibility$_{w(t-2)}$	-0.904***	-1.119***	-0.701***	-0.937***	
	(0.243)	(0.217)	(0.164)	(0.218)	
m1			-1.71*	-1.84**	-1.68*
m2			-0.08	-0.34	0.06
Hansen J Statistic (d.f.)			0.009	0.289	0.087
			(10)	(28)	(21)
No. of Groups	33	33	33	33	33
Estimation Period	1992:2005	1992:2005	1992:2005	1992:2005	1992:2005
No. of obs.	396	396	396	396	396

Hong Kong export price index is used to deflate the nominal dollar value of the ordinary exports. Columns 1-4 report estimation results of the fully specified model, while column 5 reports estimation results of the hypothetical model, which excludes $RER_{w,(t-j)}$ by assumption. As in Table 3, both GMM1 and GMM2 are one-step system GMM estimates that are obtained by estimating a system of the differenced equations and the levels equations of the model. While GMM1 assumes that RER_{ci} and RER_{w} are exogenous, GMM2 assumes that they are predetermined in the model. 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. Hansen J statistic is the test for over-identifying restrictions, asymptotically distributed as χ^2 under the null of instrument validity. Asymptotic standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses. Significance tests: *** p<0.01, ** p<0.05, * p<0.1.

Table 6
Estimates of long-run export demand elasticities

Specifications and estimations methods	Income of importing country $(\hat{\beta})$	Bilateral RMB real exchange rate $(\hat{\xi})$	Intra-regional RER flexibility $(\hat{\psi})$
Panel A: China's processing exports to 33 countries—1992-2005			
(1) Based on within estimates of the dynamic regression, Eq. 3.1	2.497*** (0.772)	−1.131*** (0.374)	−4.503*** (0.572)
(2) Based on GMM system estimates of the dynamic regression, Eq. 3.1 (both the RER vars are treated strictly exogenous)	1.011*** (0.153)	−0.257 (0.321)	−5.432*** (1.002)
(3) Based on GMM system estimates of the dynamic regression, Eq. 3.1 (both the RER vars are treated predetermined)	0.975*** (0.180)	−0.934** (0.481)	−8.465** (4.558)
Panel B: China's ordinary exports to 33 countries—1992-2005			
(1) Based on within estimates of the dynamic regression, Eq. 3.1	0.219 (0.727)	−1.175*** (0.338)	−6.022*** (0.696)
(2) Based on GMM system estimates of the dynamic regression, Eq. 3.1 (both the RER vars are treated strictly exogenous)	1.064*** (0.103)	1.049** (0.397)	−8.376 (5.743)
(3) Based on GMM system estimates of the dynamic regression, Eq. 3.1 (both the RER vars are treated predetermined)	0.946*** (0.269)	−0.715 (1.607)	−15.389 (31.130)

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

Long-run estimates (3) in both panels (A) and (B) are based on consistent system GMM estimates, whereas the estimates (1) and (2) are reported for comparison. For the dynamic estimates in each case, refer to columns (2-4) in Tables 4 and 5 respectively. Standard errors, obtained by using delta method, are reported in parentheses.

Table 7
Trade effect of a fixed exchange rate system in East Asia
(Estimates are shown for China's processing exports to 33 countries)

Importer IDs	Importing country	$\bar{y}_i$ (in million USD) [1]	$\bar{y}_i^*$ (in million USD) [2]	$\bar{y}_i^* - \bar{y}_i$ (in million USD) [3]	t-ratio [4]	$(\bar{y}_i / \bar{y}_i^*)$ [5]
101	Argentina	179	243	63.6	2.4	0.74
102	Australia	2112	2508	396.1	3.7	0.84
103	Austria	241	396	154.2	5.6	0.61
104	Belgium	1520	1691	170.7	2.2	0.90
105	Brazil	692	769	76.3	1.1	0.90
106	Canada	2183	2651	468.3	2.7	0.82
107	Denmark	511	669	158.0	4.8	0.76
108	Finland	749	785	36.7	0.7	0.95
109	France	2812	3912	1099.7	4.0	0.72
110	Germany,FR	8065	9809	1744.0	3.3	0.82
111	Greece	235	321	86.8	2.8	0.73
112	Hong Kong	42306	48473	6167.8	1.7	0.87
113	Iceland	8	13	5.1	1.9	0.61
114	Indonesia	793	844	51.3	0.9	0.94
115	Ireland	712	667	-45.2	-0.7	1.07
116	Italy	1599	2290	690.8	9.0	0.70
117	Japan	27891	42850	14958.8	8.8	0.65
118	Korea Rep	7241	9439	2198.8	7.2	0.77
119	Luxembourg	326	161	-164.8	-1.6	2.02
120	Malaysia	2189	2388	199.6	1.9	0.92
121	Mexico	972	965	-7.6	-0.1	1.01
122	Netherlands	6175	5487	-687.8	-1.0	1.13
123	New Zealand	219	273	54.3	4.3	0.80
124	Philippines	918	878	-40.0	-0.5	1.05
125	Portugal	110	156	46.2	5.4	0.70
126	Russia	636	932	296.7	5.0	0.68
127	Singapore	4565	5313	748.3	2.8	0.86
128	Spain	997	1384	386.4	6.4	0.72
129	Sweden	470	722	252.4	6.6	0.65
130	Taiwan pro	4075	6533	2457.5	10.5	0.62
131	Thailand	1397	1472	75.5	1.2	0.95
132	United Kingdom	4726	6084	1358.7	6.1	0.78
133	United States	46779	52286	5507.8	1.8	0.89
	Overall	174400	213365	38965	2.35	0.81

For each cross-section i, $\bar{y}_i$ and $\bar{y}_i^*$ represents respectively the actual and the potential volumes of China's processing exports. The potential volume is estimated under the assumption that China shares a fixed exchange rate system with the rest of East Asia and that the regional currency is perfectly flexible vis-à-vis the rest of the world. Both $\bar{y}_i$ and $\bar{y}_i^*$ are respective averages over the 1994-2005 period. The t-ratios are based on nonparametric bootstrap method.

Table 8

Estimation results of the fully specified model under alternative variable definitions and w/wo control for the supply side effect

	1	2	3	4	5	6	7	8	9	10
Dependent variable: Bilateral real exports	Processing	Processing	Ordinary	Ordinary	Processing	Processing	Processing	Ordinary	Ordinary	Ordinary
Independent variables										
y_{it-1}	0.802***	0.781***	0.570***	0.547***	0.695***	0.719***	0.741***	0.413***	0.442***	0.461***
	(0.071)	(0.073)	(0.112)	(0.109)	(0.075)	(0.087)	(0.088)	(0.093)	(0.100)	(0.104)
y_{it-2}	0.013	0.042	0.362***	0.378***	0.123	0.129*	0.099	0.387***	0.405***	0.435***
	(0.057)	(0.059)	(0.103)	(0.106)	(0.073)	(0.070)	(0.073)	(0.107)	(0.106)	(0.109)
GDP_{it}	2.647***	2.561***	1.571**	1.581**	2.849***	2.946***	2.851***	2.089***	2.229***	2.225***
	(0.521)	(0.511)	(0.626)	(0.631)	(0.519)	(0.540)	(0.543)	(0.627)	(0.625)	(0.676)
RER_{ci}	-0.749***	-0.751***	-0.903***	-0.886***	-0.659***	-0.606***	-0.665***	-0.698***	-0.646***	-0.691***
	(0.184)	(0.182)	(0.150)	(0.149)	(0.210)	(0.220)	(0.228)	(0.166)	(0.171)	(0.180)
RER_{w}	-1.082***	-1.325***	-0.289	-0.663*	-2.280***	-2.073***	-1.957***	-1.774***	-1.724***	-2.019***
	(0.378)	(0.371)	(0.301)	(0.329)	(0.548)	(0.424)	(0.483)	(0.262)	(0.288)	(0.354)
Proxy for supply shift effect	No	No	No	No	FDI stock	GDP of exporter	Gross fixed C. F.	FDI stock	GDP of exporter	Gross fixed C. F.
Deflator used	USCPI	USMPI	USCPI	USMPI	HKXPI	HKXPI	HKXPI	HKXPI	HKXPI	HKXPI
Long-run effect of RER_{w} $(\hat{\psi})$	-5.376**	-8.962**	Not	Not	-14.175*	-16.474	-13.933*	Not	Not	Not
	(2.986)	(4.646)	Significant	Significant	(7.912)	(11.173)	(8.846)	Significant	Significant	Significant
m1	-2.94***	-2.92***	-1.94*	-1.83*	-2.99***	-3.02***	-3.06***	-1.60	-1.61*	-1.56
m2	-0.07	-0.17	0.26	-0.24	-0.93	-0.72	-0.56	-0.45	-0.37	-0.44
Hansen J Statistic (d.f.)	0.588	0.585	0.318	0.380	0.501	0.712	0.484	0.794	0.474	0.524
	(28)	(28)	(28)	(28)	(26)	(27)	(27)	(27)	(27)	(27)
No. of Groups	33	33	33	33	33	33	33	33	33	33
Estimation Period	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005	1992: 2005
No. of obs.	396	396	396	396	396	396	396	396	396	396

The table reports the one-step GMM estimators of the dynamic panel model (3.1). A stacked system of the differenced equations and the levels equations of the model is estimated in every case and an extended instrument matrix is used to obtain the estimators. Both RER_{ci} and RER_{w} are treated to be predetermined (weakly exogenous) in the model. When the model is augmented by including proxy for the supply shift effect, only the lagged (t-2) term of proxy variables is used in the estimation. This was necessary because these proxies are too endogenous. The long-run coefficients of other variables are omitted for reporting convenience. See footnotes to Table 4 for interpretations of m1, m2 and Hansen J statistics. Standard errors, asymptotically robust to cross-section and time-series heteroscedasticity, are reported in parentheses. Significance tests: *** p<0.01, ** p<0.05, * p<0.1