ON TRADE BALANCE AND EXCHANGE RATES: FURTHER EVIDENCE FROM CHINA'S BILATERAL TRADE*

1. Introduction

Since its economic liberalization of early 1980s, China has enjoyed a large and growing surplus in her foreign trade balance. The trade surplus increased from 1% of GDP in 1982 to 2% in 1990 and 7% in 2008. As a result, a number of critics have pointed out that China's large and growing trade surplus reflects primarily the government's deliberate policies to keep its currency (renminbi) undervalued; and an effective solution to the large current account surplus requires major revaluations of real renminbi[1]. This recommendation is consistent with long-run predictions of conventional trade theory. Such a revaluation is expected to increase the relative price of Chinese exports and contribute to a decline in its trade surplus[2].

However, as Fawley and Juvenal (2011) and Holtz-Eakin (2003) have pointed out, two factors may significantly hinder the effect of renminbi revaluation on its trade balance. First, China has served as the manufacturer of the world and a place of assembly for the Asian region due to its low labor costs. Intermediate goods are exported to China, which are then assembled and exported to the rest of the world. Given that more than 70% of these exports originate beyond Chinese borders, most of the value added is set in

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[1] See for example, Frank James's NPR report "Fed Up! Bernanke (Gently) Takes on Critics" at: http://www.npr.org/blogs/itsallpolitics/2010/11/19/131441539/fed-up-bernanke-gently-takes-on-critics.

[2] The short-run consequences of such revaluation are, however, not clearly understood. The trade surplus might further increase immediately after the revaluation due to the (inverted) J-curve effect. As Krueger (1993) and others have pointed out, the volume of imports and exports in the short run is determined in part by pre-existing contracts which must be honored. As such, a currency revaluation will improve the trade balance by primarily reducing the price and thus total spending on imports.

other countries and in terms of other currencies. Thus only up to 30% of the value added for Chinese exports is subject to the effects of renminbi revaluation. Following a renminbi revaluation, Chinese manufacturers can import intermediate imports at lower costs and pass the benefits to final prices, largely offsetting the effect of revaluation. Second, a country's trade balance is partly a reflection of its domestic preferences for saving and investment, which do not necessarily depend on the value of its exchange rate. In that sense, revaluing renminbi may not dramatically affect Chinese preferences for domestic saving and investment, and may have a minimal effect on its trade balance. As such, economic theory does not provide a clear assessment on the effect of renminbi on Chinese foreign trade. Thus, this question must be addressed empirically.

In this paper, we will examine the relationship between China's bilateral trade and real exchange rates with annual data for thirteen trading partners over 1981-2008. In particular, we attempt to address a number of important but yet unanswered questions as follow: (1) Is there an evidence of significant regime shift in Chinese bilateral trade balance and real exchange rates; and if so, is the timing of break consistent with the shift in the exchange rate regime in 1994 and/or 2005? (2) Accounting for domestic and foreign incomes, is there a long-run relationship between the real exchange rates and the trade balance; and if so, is the relationship subject to regime shifts? (3) Are adjustments of the trade balance to deviations from equilibrium symmetric, consistent with competitive markets, or asymmetric suggesting the possibility of bottlenecks in China's trade with its partners? (4) What are the causal relations between real exchange rates and the bilateral trade balance, and how does the trade balance respond to various exogenous shocks? Finally, (5) what are the sources of long-run variations in China's bilateral trade balance?

We address the first question by applying two tests of unit roots with endogenous breaks, Lee-Strazicich (2003) and Zivot-Andrew (1992). For questions two and three, we apply four alternative tests of cointegration – Johansen's (1991) maximum likelihood, Gregory-Hansen's (1996) test of cointegration with structural break, Enders and Siklos's (2001) residual-based momentum autoregressive (MTAR) model, and Pedroni's (1999, 2001, 2004) panel cointegration tests. Finally, we employ tests of causality, impulse response functions and variance decompositions to address questions four and five.

The remainder of this paper is organized as follows: Section 2 provides a brief review of the literature. Section 3 describes the empirical model. Section 4 explains the econometric methods.

Section 5 explains the data and construction of the variables. Section 6 reports the empirical results and Section 7 provides a summary of the results and main conclusions.

2. Previous literature

A large and growing body of literature has examined the responsiveness of the trade balance to exchange rates. Several of these studies have focused on China. Table 1 provides a summary of this literature. Bahmani-Oskooee and Ratha (2004) group these studies into two categories based on the level of data aggregation. One category employs aggregate data. Brada *et al*. (1993), Zhang (1998, 1999a, 1999b), Wei (1999), Wang (2007), Garcia-Herrero and Koivu (2009), Zhang and Sato (2009) and Yaya and Lu (2010) belong to this category. Several of these studies investigate the long-run effect of the exchange rate on the trade balance. Brada *et al*. (1993) find a long-run relation between real RMB and trade balance with quarterly data over 1980-1989 using Johansen's procedure. Zhang (1999a) reports similar results with monthly data over 1991-1996. However, the long-run effect of real effective exchange rates is not sizable. Zhang (1999b) reports the robustness of these findings to extension of the sample to 1986-1997. Wei (1999) finds no evidence of cointegration in a model which additionally includes domestic and foreign money supplies and is estimated using the Engle-Granger's (1987) procedure. In contrast, Wang (2007) finds a stable relation between real effective exchange rates and trade balance in quarterly data over 1997-2006. Garcia-Herrero and Koivu (2009) report similar findings with monthly data over 1994-2005. However, a real appreciation tends to reduce both China's exports and imports, thus limiting the net impact of exchange rate policy on the trade surplus.

A number of other studies have examined the dynamics of trade-balance exchange rate relations using tests of Granger-causality, impulse response functions and variance decompositions. Zhang and Sato (2009) find limited evidence on short-run relations between the two variables using quarterly data over 1987-2007. Zhang (1998) and Yaya and Lu (2010) find that the trade balance Granger-causes real effective exchange rates but not the vice versa.

Bahmani-Oskooee and Brooks (1999) have raised the possibility that the above findings might be subject to aggregation bias. A country's trade balance and/or exchange rate may improve with respect to one trading partner while deteriorating with respect

TABLE 1 - *Previous Literature*

AUTHOR (YEAR)	METHODOLOGY	FREQUENCY/ PERIOD	DEPENDENT VARIABLE	EXPLANATORY VARIABLES	COUNTRY	FINDINGS
Bahmani-Oskooee and Wang (2008)	ARDL, ECM	A:1978-2002	ln(IM/EX)	lnY, lnY_c, RER	USA	Long-run effect of RER positive and significant in 34 cases, with 22 cases of support for J-curve.
Bahmani-Oskooee and Wang (2006)	ARDL, ECM, J.J. C.	Q:1983-2002	ln(IM/EX)	lnY, lnY_c, RER	13 partners	Depreciation has significant short-run effect but disappears in the long run.
Brada *et al*. (1993)	ADF, J.J. C.	Q:1980 -1989	ln(TB)	lnY, lnY_c REER	Total	RER has a long-run effect.
Garcia-Herrero and Koivu (2009)	ADF, J.J. C.	Q:1994-2005	EX, IM	lnY, lnY_c REER	Total/13 partners	Appreciation reduces both exports and imports, thus net impact is limited.
Narayan (2006)	ARDL	M:1979-2002	ln(EX/IM)	RER	USA	Devaluation improves TB but no J-curve effect.
Wang (2007)	ADF, J.J. C., VECM	Q:1997-2006	lnTB	lnY, lnY_c REER	Total	Devaluation increases TB but the effect is less than income.
Wei (1999)	ADF, E.G. C.	M:1986-1996	EX/IM	MS_c/MS, Y_c/Y	Total	Lack of long-run relationship between trade balance and its determinants.
Yaya & Lu (2010)	Causality	M:1994-2009	TB	REER	Total	TB Granger causes REER but not vice versa.
Zhang (1998)	Granger causality	M:1991-1996	TB	Y, Y_c, RER	Total	TB Granger causes RER but not vice versa.
Zhang (1999a)	ADF, J.J. C., ECM	M:1991-1996	TB	Y, Y_c, RER	Total	The long-run effect of RER on TB is not sizable; no J-curve effect.
Zhang (1999b)	ADF, J.J. C.	M:1986-1997	TB	Y, Y_c, REER	Total	The effect of REER on TB is moderate; Y_c and Yare more important; no J-curve.
Zhang and Sato (2009)	Structural VAR, VD	Q:1987-2007	TB	$Y_{u.s}$, Y_c, RER	Total	Dynamic effect of RER on TB very limited.

Notes: ADF is augmented Dickey-Fuller; ARDL is autoregressive distributed lag model; ECM is error correction model; E.G. C. is Engle Granger's test of cointegration; VECM is vector error correction model; VAR is vector autoregressive model; VD is variance decomposition; J.J. C. is Johansen and Juselius's test of cointegration; M, Q and A refer to monthly, quarterly and annual frequencies; RER and REER are real and real effective exchange rates; Y, Y_c and $Y_{u.s.}$ are foreign and Chinese, and the U.S. real incomes; TB is trade balance; EX and IM are real exports and imports; MS_c and MS are stocks of domestic and foreign money.

to others. In response, a number of studies have examined the relationship with disaggregated data. Narayan (2006) and Bahmani-Oskooee and Wang (2006, 2008) fall into this category. Narayan (2006) investigates the responsiveness of China's trade with the U.S. to changes in exchange rates using the autoregressive distributed lag (ARDL) model. Both long-run and short-run exchange rate elasticities are found positive and significant, suggesting that a real depreciation of RMB will improve China's trade balance in both long-run and short-run. Bahmani-Oskooee and Wang (2006) find significant short-run effects of a real depreciation on China's trade balance, which disappear in the long run in most cases. Also, using commodity-level disaggregated data (Bahmani-Oskooee and Wang, 2008) provides a weaker evidence of real exchange rate-trade balance relationship.

Viewed as a whole, available evidence on the effect of exchange rates on China's trade balance is mixed in both long run and short run, and sensitive to the choice of the empirical model, sample period, and estimation technique. In fact, the sensitivity of the results to sample periods may suggest the possibility of structural breaks or inherent nonlinearities in the relationship.

3. Empirical model

The specification of the trade balance model follows Bahmani-Oskooee and Brooks (1999) and others[3],

$$TB_{j,t} = \alpha + \beta Y_{d,t} + \gamma Y_{j,t} + \delta Q_{j,t} + e_{j,t} \quad (1)$$

where TB_j is China's trade balance with the j'th trading partner; t is time; Y_j is her real income; Y_j is real income for the j'th trading partner; Q_j is the bilateral real exchange rate with the j'th trading partner; and e_j is a zero mean error term. We measure TB_j by logarithm of China's exports/imports ratio with the j'th trading partner. Thus values greater than zero imply a trade surplus. We measure Q_j by logarithm of the price of Chinese goods relative to that of j'th trading partner. Thus, a rise in Q_j reflects a real appreciation of renminbi with respect to the trading partner's currency. Finally, we measure Y_d and Y_j by logarithms of real GDP in each country. Conventional trade theory suggests that a rise in domestic income increases imports and reduces the trade surplus ($\beta < 0$); a rise in partner's income increases exports and improves the trade surplus

[3] See for example references in Table 1.

($\gamma < 0$) and a rise in Q_j implies a rise in relative price of exports, and a fall in exports relative to imports ($\delta < 0$).

4. Empirical methods

The empirical work carried out in this study requires variables that are non-stationary in level but stationary in first differences. Thus, we begin with tests of unit roots using the traditional augmented Dickey-Fuller (ADF) procedure. A potential shortcoming of the traditional ADF test is its inability to capture structural breaks in the time series. However, the importance of allowing for structural breaks in unit root tests is now well documented. This is potentially a serious shortcoming given the trade and exchange rate adjustments in China over the sample period. To address this issue, we employ two alternative unit root tests with endogenous structural breaks, namely Lee and Strazicich (2003) and Zivot and Andrews (1992). Next, we examine the existence of long-run relations among variables using four alternative tests of cointegration. We begin with the well-known maximum-likelihood procedure proposed by Johansen and Juselius (1990) and Johansen (1991). However, this test does not account for possible endogenous breaks as well as nonlinearities in the relationship. Also, the small-sample properties of this procedure are not clearly understood.

We assess the possibility of structural breaks in the long-run relations by applying Gregory-Hansen's (1996) residual-based test of cointegration with endogenous regime shifts. Similarly, we allow for asymmetric adjustments to deviations from equilibrium by applying the momentum threshold autoregressive (MTAR) model of Enders and Siklos (2001). This procedure involves estimating the following model:

$$\Delta\hat{e}_t = \rho_1 M_t \hat{e}_{t-1} + \rho_2 (1 - M_t)\hat{e}_{t-1} + \sum_{i=1}^{n} \gamma_i \Delta\hat{e}_{t-i} + u_t \qquad (2)$$

where $\hat{e}_t$ are estimates of equilibrium error in equation (1), $u_t \sim IID(0,\sigma^2)$, and the Heaviside indicator, M_t, is set according to:

$$M_t = 1 \quad if\ \Delta\hat{e}_{t-1} \geq \tau$$
$$M_t = 0 \quad if\ \Delta\hat{e}_{t-1} < \tau$$

Here the threshold, τ, is set endogenously using Chan's (1993) methodology. The MTAR model allows the residuals to exhibit different degrees of autoregressive decay depending on the momentum in lagged residuals. Test statistics Φ^*_μ are used to examine the

null hypothesis of no-cointegration (Ho: $\rho_1 = \rho_2 = 0$). Critical values are reported in Wane *et al*. (2004). If the null hypothesis of no-cointegration is rejected, then one may proceed to test for the null hypothesis of symmetry (Ho: $\rho_1 = \rho_2$). If the null hypothesis of symmetry is rejected then the model exhibits more decay in one direction than the other.

Finally, we address the small sample problem inherent in our time series analysis by testing for a long-run relation in equation (1) using Pedroni's (1999, 2001, 2004) tests of panel cointegration. This technique has three additional advantages: it allows the short-run dynamics to vary among members of the panel; it allows for heterogeneous unobservable effects and deterministic trends; and it allows for different cointegrating vectors for each cross section. Pedroni (1999, 2001, 2004) proposes seven test statistics, four of which are based on within-dimension and three of which are based on between-dimension.

Tests of cointegration, while consistent with long-run relations, do not provide information on the dynamics of the relationship between the trade balance and exchange rates. We address these issues using tests of Granger-causality, impulse response functions and variance decompositions using the VAR methodology of Toda and Yamamoto (1995). This procedure is appealing in three respects: it is robust to the underlying stochastic properties of individual variables; it does not depend on the cointegration properties of the variables in the model; and the VAR model can be estimated in levels.

Application of Toda-Yamamoto procedure involves the following three steps: first, we specify a VAR model in terms of the level of the variables and determine the optimal lag length (k) using a lag-order selection criterion. Next, we augment lag k by adding the maximal order of integration d_{max}, and estimate the VAR($k+ d_{max}$) using OLS. For the two variables of our primary interest, trade balance (TB) and real exchange rates (Q), the equations are,

$$TB_{j,t} = \sum_{i=1}^{k+d_{\max}} \alpha_i TB_{j,t-i} + \sum_{i=1}^{k+d_{\max}} \beta_i Y_{d,t-i} + \sum_{i=1}^{k+d_{\max}} \gamma_i Y_{j,t-i} + \sum_{i=1}^{k+d_{\max}} \delta_i Q_{j,t-i} + e_{1,t} \qquad (3a)$$

$$Q_{j,t} = \sum_{i=1}^{k+d_{\max}} \alpha_i' TB_{j,t-i} + \sum_{i=1}^{k+d_{\max}} \beta_i' Y_{d,t-i} + \sum_{i=1}^{k+d_{\max}} \gamma_i' Y_{j,t-i} + \sum_{i=1}^{k+d_{\max}} \delta_i' Q_{j,t-i} + e_{2,t}. \qquad (3b)$$

Tests of Granger-causality is conducted as follows: the null hypothesis of no-causality from Q to TB is carried out by testing the

joint significance of δ_i ($i=1,\ldots, k$) on the first k lags of Q in equation (3a). Similarly, the null hypothesis of no-causality from TB to Q is examined by testing the joint significance of $\alpha_i^{'}$ ($i=1,\ldots, k$) on the first k lags of TB in equation (3b).

Finally, we examine the dynamics of the relationship between the trade balance and real exchange rates using impulse response functions, and evaluate the source of variations in their forecast errors using variance decompositions. Impulse response functions for TB and Q provide information on the transitory versus permanent nature of shocks to the system; and variance decompositions measure the variability of forecast errors in response to different innovations in the system.

5. Data and variables

The empirical work in this study is carried out using bilateral trade data between China and 13 of her major trading partners as well as measures of their real income and bilateral real exchange rates. We use annual data for 1981-2008. The trading partners are Australia, France, Germany, Hong Kong, India, Italy, Japan, Malaysia, the Netherlands, Saudi Arabia, Singapore, United Kingdom and United States. The trade balance (TB_j) is measured as the logarithm of the ratio of China's exports to and imports from trade partner j. The raw data for exports and imports are obtained from various issues of China's Statistical Yearbook. Real income for China (Y_c) and the j'th trade partner (Y_j) are measured by their logarithms of real GDP in U.S. dollars. Real exchange rate (Q_j) is the logarithm of the bilateral real exchange rate between renminbi and currency of the j'th trade partner. It is constructed as the logarithm of $(P_c*E_j)/P_j$, where P_c and P_j are CPI price indices for China and j'th trade partner, and E_j is the nominal bilateral exchange rate, measured as the number of trading partner j's currency per unit of renminbi. Data for real GDP, nominal exchange rates and price levels are collected from International Financial Statistics, IMF.

Table 2 provides summary statistics on China's trade volumes and trade shares *vis-à-vis* each trade partner. At the top of the list are the three countries of U.S., Japan and Hong Kong with 13.2%, 10.40%, and 7.94% of shares of trade. For the remaining 10 partners, the trade shares range from a high of 4.49% for Germany to low of 1.49% for Italy.

TABLE 2 - *China's Shares of Trade with Major Partners (2008)*

PARTNER	% EXPORTS	% IMPORTS	% TRADE
Australia	1.55	3.31	2.33
France	1.64	1.38	1.52
Germany	4.14	4.93	4.49
Hong Kong	13.33	1.14	7.94
India	2.21	1.79	2.02
Italy	1.86	1.03	1.49
Japan	8.12	13.3	10.41
Malaysia	1.50	2.83	2.09
Netherland	3.21	0.47	2.00
Saudi Arabia	0.76	2.74	1.63
Singapore	2.26	1.78	2.05
UK	2.52	0.84	1.78
USA	17.64	7.18	13.02
Total	60.74	42.72	52.77

Notes: % Exports is exports to each trade partner as a percentage of total China's exports; % Imports is China's imports from each trade partner as a percentage of total imports from the rest of world; % Trade is trade volume with each partner as a percentage of the total trade volume.

Source: China Statistical Yearbook, 2009.

6. EMPIRICAL RESULTS

We begin with traditional ADF tests of unit roots, which are reported in Table 3. With few exceptions, the test statistics fail to reject the null hypothesis of unit roots in levels but reject the hypothesis with first-differenced data at the conventional 5% significance level. Exceptions are levels of trade balance for India, income for Japan and real exchange rate for the Netherlands, where the null hypothesis is not rejected at the 1% level; and changes in income for France, Hong Kong, Italy and Singapore where the null hypothesis is rejected at the 10% level.

Next, we explore the robustness of unit root tests to possible structural breaks by applying Lee-Strzicich and Zivot-Andrews procedures. The results of these tests are reported in Table 4. Two interesting patterns emerge. First, allowing for structural breaks weakens the evidence in favor of non-stationarity in variables. Now, the null hypothesis of unit roots is rejected in five countries for trade balance, in two countries for exchange rates, and in four countries for foreign income. Second, most structural breaks in the trade balance and exchange rates occur in mid 1990s, which coincide with

TABLE 3 - *Augmented Dickey-Fuller Tests of Unit Roots*

	TB		Y_j		*Q*	
	C	C/T	C	C/T	C	C/T
			LEVELS			
Australia	-0.85	-0.86	0.45	-2.17	-2.47	-2.28
France	-1.78	-3.95*	-0.76	-5.09*	-2.55	-2.03
Germany	-1.30	-2.23	-0.29	-4.03*	-2.22	-2.02
Hong Kong	0.47	-3.31	0.77	-3.19	-2.41	-0.26
India	-3.23*	-3.94*	-1.86	0.67	-3.58*	-3.54
Italy	-0.49	-2.63	-2.32	-0.42	-2.68	-1.93
Japan	-2.21	-2.22	-3.38*	-4.08*	-3.00*	-1.89
Malaysia	-2.65	-2.74	-0.49	-1.83	-2.67	-2.64
Netherland	-0.005	-3.33	0.07	-1.96	-4.16*	-4.66*
Saudi Arabia	-1.67	-2.83	0.68	-5.64*	-2.26	-2.67
Singapore	-1.99	-1.66	-0.38	-2.03	-2.74	-1.83
UK	-1.05	-2.79	-0.07	-2.96	-1.94	-1.60
USA	-1.05	-3.50	-2.51	-3.14	-3.54*	-2.07
			FIRST DIFFERENCES			
Australia	-4.53*	-5.31*	-3.88*	-4.01*	-5.20*	-5.32*
France	-10.82*	12.09*	-2.86	-2.82	-4.88*	-5.13*
Germany	-4.98*	-5.02*	-3.79*	-3.69*	-5.15*	-5.24*
Hong Kong	-3.26*	-3.71*	-2.82	-2.65	-3.71*	-4.71*
India	-6.22*	-6.26*	-1.38	-3.88*	-3.92*	-3.90*
Italy	-5.47*	-5.76*	-2.74	-3.51	-4.74*	-5.17*
Japan	-4.68*	-4.46*	-7.81*	-7.66*	-3.55*	-4.68*
Malaysia	-4.66*	-4.76*	-4.19*	-4.11*	-4.66*	-4.82*
Netherland	-4.88*	-5.23*	-3.88*	-3.78*	-8.10*	-7.98*
Saudi Arabia	-5.42*	-4.17*	-3.90*	-4.06*	-4.26*	-4.68*
Singapore	-5.37*	-7.08*	-2.95	-2.88	-4.54*	-4.90*
UK	-4.67*	-4.40*	-3.68*	-3.44	-4.41*	-4.96*
USA	-7.55*	-7.46*	-4.63*	-5.14*	-3.97*	-4.64*

Notes: the table reports student t-statistics for tests of unit roots; optimal lag lengths selected by Schwarz Information Criterion; 5% significant level is indicated by '*'. The variables are trade balance (*TB*), *j*'th trading partner's real income (Y_j) and real exchange rate (*Q*). Models include an intercept (C), or intercept and trend (C/T).

TABLE 4 - *Student t-Statistics for Tests of Unit Roots with Endogenous Breaks (1981-2008)*

	Lee-Strazicich			Zivot-Andrews		
	TB	Y_j	*Q*	*TB*	Y_j	*Q*
Australia	-1.976 (1983)	-2.450 (1984)	-2.475 (1997)	-3.165 (1990)	-2.923 (2004)	-3.497 (1998)
France	-4.115* (1984)	-3.503 (1993)	-1.862 (1994)	-3.624 (1989)	-3.772 (1988)	-2.984 (1990)
Germany	-4.116* (1987)	-3.018 (1989)	-2.298 (1994)	-4.679 (2001)	-3.177 (1997)	-3.247 (1990)
Hong Kong	-2.171 (1996)	-3.707* (1987)	-1.338 (1993)	-4.211 (1994)	-5.249* (1998)	-3.236 (1990)
India	-2.513 (1992)	-1.436 (1984)	-3.178 (1995)	-3.500 (1991)	-2.301 (2003)	-4.044 (1996)
Italy	-2.835 (1984)	-2.391 (1984)	-1.709 (1989)	-3.840 (1993)	-2.170 (1987)	-3.061 (1990)
Japan	-4.617* (1989)	-3.237 (1985)	-2.483 (1989)	-5.204* (1990)	-4.050 (2002)	-2.968 (1990)
Malaysia	-3.525 (1989)	-3.089 (1998)	-1.955 (1997)	-3.458 (1987)	-6.083* (1998)	-4.771 (1990)
Netherland	-2.224 (1984)	-3.239 (1989)	-3.892* (1985)	-6.249* (1995)	-3.231 (1997)	-5.634* (2002)
Saudi Arabia	-3.377 (1991)	-1.296 (1986)	-3.127 (1983)	-3.305 (1989)	-3.555 (2003)	-4.228 (1990)
Singapore	-2.347 (1987)	-2.897 (1998)	-1.611 (1993)	-2.517 (1988)	-3.951 (2001)	-5.151* (1990)
UK	-2.210 (1986)	-3.285 (1989)	-3.121 (1993)	-5.089* (1987)	-4.974* (1993)	-3.032 (1990)
USA	-3.027 (1993)	-4.701* (1996)	-1.979 (1983)	-4.610 (1994)	-2.702 (1997)	-4.121 (1990)

Notes: *TB* is the logarithm of exports to imports ratio; Y_f is the ln of real GDP; *Q* is ln real exchange rate. The values in parentheses are dates of the structural break. The 5% critical values for the unit root tests are -3.566 for Lee-Strazicich and -4.800 for Zivot-Andrews. Rejection of unit roots hypothesis is denoted by *.

China's change in the exchange rate regime. Given these findings, we proceed to tests of cointegration, bearing in mind that support for cointegration, if any, may partially reflect the stationary nature of some variables in the model.

Table 5 reports the results of tests of cointegration using the maximum-likelihood procedure. Following Cheung and Lai (1993),

TABLE 5 - *Maximum Likelihood Tests of Cointegration*

	AUSTRALIA	FRANCE	GERMANY	HONG KONG	INDIA	ITALY	JAPAN
Max-Eigen test							
r = 0	15.66	34.60*	22.45	16.38	31.59*	22.88	20.17
r ≤1	7.76	18.47	12.92	11.37	12.50	9.71	8.41
r ≤2	6.23	8.91	5.65	7.54	5.56	7.09	6.78
r ≤3	0.09	0.02	3.66	0.23	0.09	0.00	2.98
Trace test							
r = 0	29.75	61.99*	44.68	35.52	49.74*	39.69	38.34
r ≤1	14.09	27.40	22.23	19.14	18.14	16.81	18.17
r ≤2	6.31	8.93	9.31	7.77	5.64	7.10	9.76
r ≤3	0.09	0.02	3.66	0.23	0.09	0.00	2.98

	MALAYSIA	NETHERLAND	SAUDI ARABIA	SINGAPORE	UK	USA	5% critical value
Max-Eigen test							
r = 0	28.99*	16.41	16.83	20.13	22.39	35.06*	27.58
r ≤1	11.04	11.39	14.61	9.31	10.77	13.59	21.13
r ≤2	4.57	7.19	9.42	4.97	6.26	12.14	14.26
r ≤3	0.35	4.79	0.21	2.34	3.46	0.002	3.84
Trace test							
r = 0	44.94	39.78	41.06	36.75	42.89	60.79*	47.86
r ≤1	15.96	23.37	24.24	16.62	20.50	25.72	29.80
r ≤2	4.92	11.97	9.63	7.31	9.73	12.14	15.49
r ≤3	0.35	4.79	0.21	2.34	3.46	0.002	3.84

Notes: r is the number of cointegrating relations; tests include 2 lags; test statistics are adjusted for small sample size by multiplying the original values by (T-NK)/T, where T is sample size, N number of variables and K number of lags; 5% significant level is indicated by '*'; the model includes an intercept but no trend.

we have adjusted the test statistics for the sample size[4]. The 5% critical values are reported in the last column of the table. Two important findings deserve attention. First, for nine-out-of-thirteen countries, the null hypothesis of no-cointegration cannot be rejected using both maximum eigenvalue and trace test statistics. Second, as for the remaining four countries – France, India, Malaysia and the USA – there is evidence of at least one cointegrating relation. However, such relation might potentially reflect the existence of one or more stationary variables in the system. As such, we conclude that evidence in favor of long-run relations between the trade balance and real exchange rates is rather weak.

Table 6 examines the robustness of cointegration tests to the possibility of structural breaks in the long-run relations using

TABLE 6 - *Gregory-Hansen's Tests of Cointegration*

	Minimum t-Statistics
Australia	-3.141 (1998)
France	-4.061 (1990)
Germany	-3.787 (1991)
Hong Kong	-3.379 (1994)
India	-3.717 (1994)
Italy	-4.236 (1991)
Japan	-4.646 (1990)
Malaysia	-3.882 (1987)
Netherland	-3.666 (1994)
Saudi Arabia	-3.887 (1991)
Singapore	-4.010 (1989)
UK	-3.854 (1994)
USA	-3.979 (1994)

Notes: dates of structural break are in parentheses. The 5% critical value for cointegration test is -5.28.

Gregory-Hansen (1996) procedure. Two findings stand out. First, there is evidence of structural breaks occurring in the 1990s in China's trade with almost all trading partners. Second, allowing

[4] This involves multiplying the original statistics by (T-NK)/T, where T is the sample size, N is the number of variables in the system and K is the lag length.

for structural breaks eliminates any evidence of long-run relations between the trade balance and real exchange rates at the 5% significance level.

Table 7 reports the results of tests of cointegration under asymmetric adjustments. Now the null hypothesis of no-cointegration cannot be rejected in China's trade with all trade partners but France. As for France, there is also evidence of significant asymmetric adjustments. In particular, the trade balance adjusts much faster to positive deviations from equilibrium.

TABLE 7 - *Tests of Cointegration and Asymmetric Adjustments (MTAR Model)*

	τ	ρ^{+}	ρ^{-}	Φ^{*}_{μ}	t	$Q_{LB}(2)$
Australia	0.006	-0.339 (0.331)	-0.199 (0.147)	1.352	0.399	0.319
France	-0.020	-1.293* (0.269)	-0.492* (0.230)	12.040*	2.591*	1.534
Germany	0.049	-0.833 (0.497)	-0.583* (-2.918)	4.863	0.502	1.148
Hong Kong	-0.005	-0.199 (0.372)	-0.372* (0.147)	3.269	-0.446	1.808
India	-0.012	-1.240* (0.424)	-0.805* (0.292)	6.331	0.982	1.837
Italy	-0.081	-0.552 (0.841)	-0.954* (0.246)	8.051	-0.519	0.368
Japan	-0.081	-0.878 (0.254)*	-0.431* (0.147)	9.257	1.600	0.218
Malaysia	0.052	-1.268* (0.398)	-0.345* (0.161)	7.069	2.182*	0.823
Netherland	-0.070	-0.511 (0.623)	-0.427* (0.166)	3.518	0.133	1.931
Saudi Arabia	0.114	-0.511 (0.312)	-0.428* (0.133)	6.444	0.247	0.425
Singapore	0.042	-0.038 (0.559)	-0.766* (0.216)	6.636	-1.331	0.101
UK	0.039	-1.029* (0.431)	-0.302 (0.195)	3.597	1.625	0.039
USA	-0.051	-0.326 (0.403)	-0.900* (0.255)	6.225	-1.363	0.119

Notes: Φ_{μ} is the test statistic for the null hypothesis that $\rho_1=\rho_2=0$; 5% critical value from Table 5 of Wane *et al*. (2004) is 11.47. $t(\rho_1=\rho_2)$ is the t statistic for the null hypothesis that the adjustment process is symmetric; standard errors of parameter estimates in parentheses; significance at 5% level is represented by *.

Table 8 reports the results of seven test statistics for panel cointegration proposed by Pedroni (1999, 2001, 2004)[5]. Now the empirical results are mixed. The null hypothesis of no-cointegration is rejected in the PP- and ADF-test statistics for both panel and group tests, but not rejected in the remaining three test statistics at the 5% level of significance. However, the findings may reflect the stationary nature of real exchange rates as suggested in tests of panel unit roots[6].

TABLE 8 - *Pedroni's Panel Cointegration Test Statistics (13 Trading Partners 1981-2008)*

Test Statistics	Values
Panel v-statistic	-0.37 (0.642)
Panel rho-statistic	-1.63 (0.052)
Panel PP-statistic	-7.10* (0.000)
Panel ADF-statistic	-3.92* (0.000)
Group rho-statistic	0.63 (0.735)
Group PP-statistic	-4.69* (0.000)
Group ADF-statistic	-3.44* (0.000)

Notes: '*' denotes significant at 5% level. Critical values are from Pedroni (1999).

Viewed as a whole, tests of cointegration do not support the existence of a long-run relationship between the trade balance and real exchange rates between China and its trade partners. As such, a revaluation of the real exchange rate might not adequately address the surplus in China's trade balance even in the long-run.

Next, we examine the dynamics of the real exchange rate-trade balance relation using a VAR model following Toda and Yamamoto (1995) procedure. Table 9 summarizes the results of tests of causality between real exchange rates and the trade balance *vis-à-vis* each trade partner. In all cases, the null hypothesis that a change in the level of real exchange rates does not Granger-cause the trade balance is not rejected. That is reflected in the statistically insignificant chi-

[5] We also performed tests of panel unit roots using Maddala and Wu's (1999) procedure. The results fail to reject the null hypothesis for foreign income and the trade balance, but reject the hypothesis for real exchange rates. Thus, the panel cointegration tests should be handled with care.

[6] The results of these tests are available upon request.

squared statistics reported in column 3. Similarly, the null hypothesis that a change in the level of trade balance does not Granger-cause real exchange rates is not rejected for all but two countries – France and Japan. As shown in column 4 of the table, only for these two countries the chi-squared statistics are significant at the 5% level.

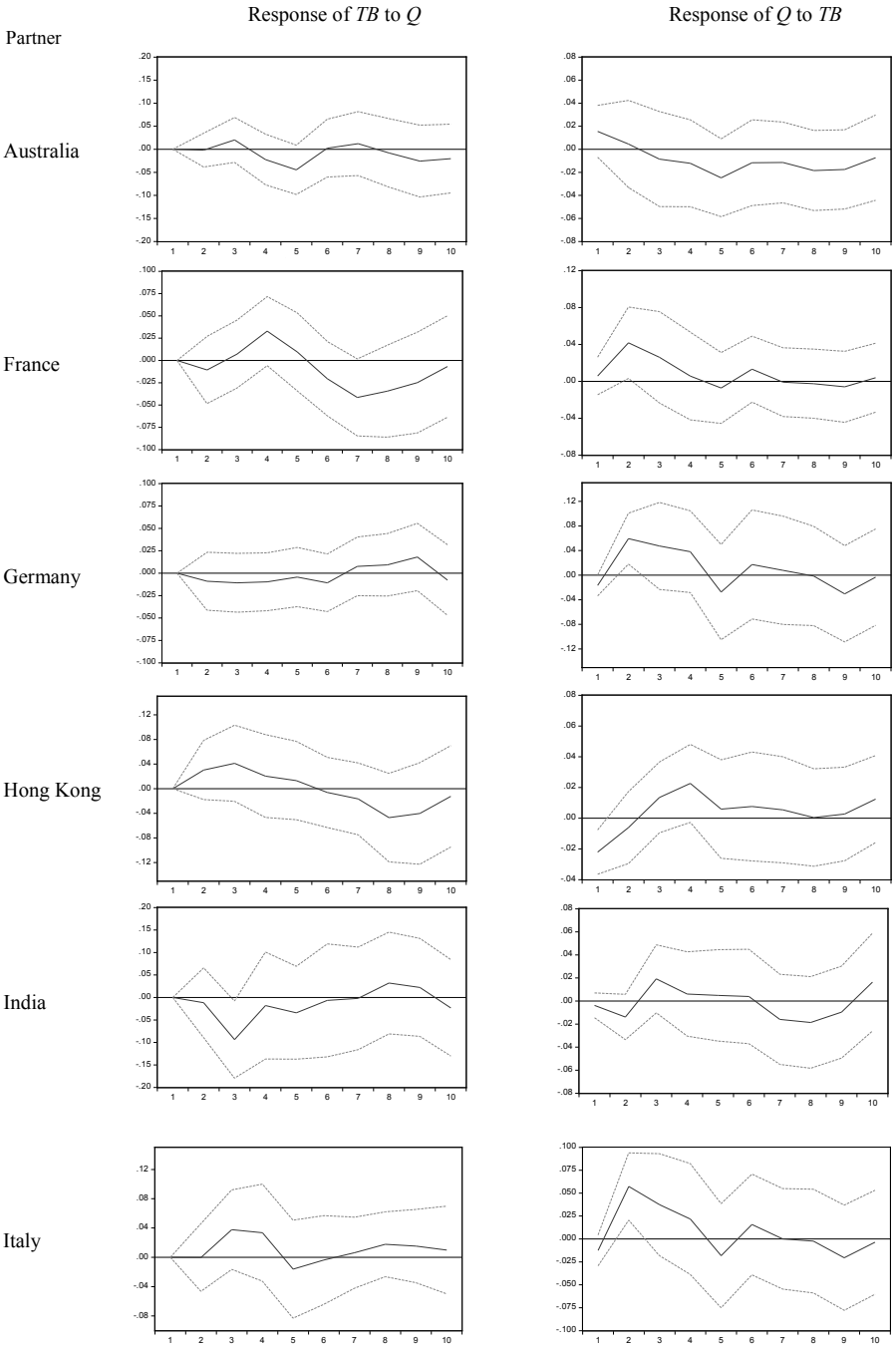
TABLE 9 - *Tests of Causality Using VAR(K*+d_{max})*

	K*	*Q≠>TB*	*TB≠>Q*
Australia	4	1.704	1.233
France	4	0.209	29.497*
Germany	4	1.172	0.299
Hong Kong	1	1.821	0.902
India	4	1.534	4.576
Italy	4	0.620	0.765
Japan	4	0.017	7.157*
Malaysia	1	0.334	2.969
Netherland	4	0.179	0.467
Saudi Arabia	4	2.376	4.295
Singapore	4	3.454	0.629
UK	4	0.675	1.302
USA	4	1.776	2.903

Notes: K* is the optimum lag length; d_{max} is the highest degree of integration in a variable in the system; F-statistics associate with tests of non-causality between real exchange rates and trade balance are reported in the last two columns; 5% significant level is indicated by '*'.

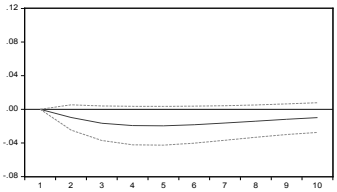
Figure 1 provides additional information on the dynamics of the trade balance-exchange rate relationship by plotting the corresponding impulse response functions along with their 95% confidence bounds. The response of the trade balance to one standard deviation shock in real exchange rates is reported in column 2. In all but one case, the confidence bound includes the zero line, suggesting that the response of trade balance to real exchange rate shocks is not significantly different from zero. The sole exception is trade with Singapore, where the trade balance improves up to five years following an exchange rate appreciation. The response of real exchange rates to shocks in the trade balance, reported in column 3, follows similar patterns. Thus, these findings reinforce the earlier results on tests of causality.

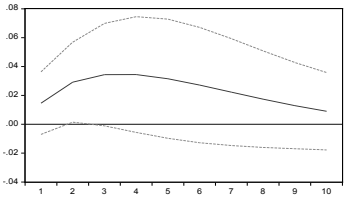
FIGURE 1 - *Impulse Response Functions for the Trade Balance (TB) and Real Exchange Rates (Q)*



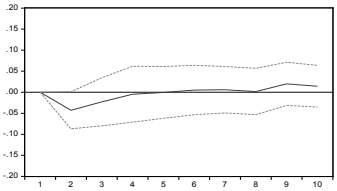
Continued FIGURE 1

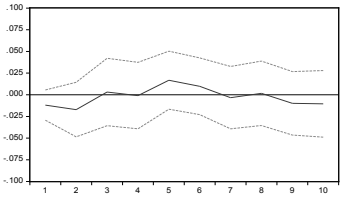
Japan



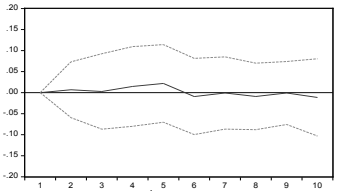


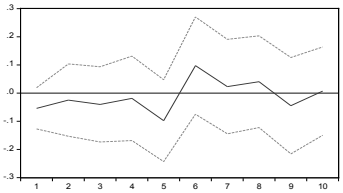
Malaysia



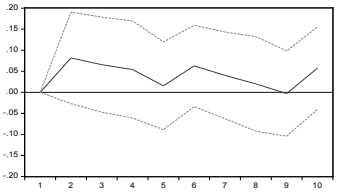


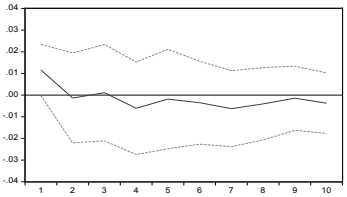
Netherland



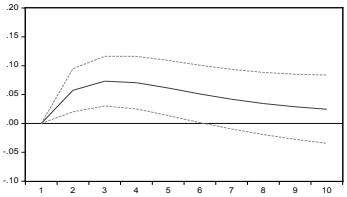


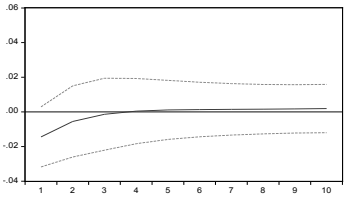
Saudi Arabia



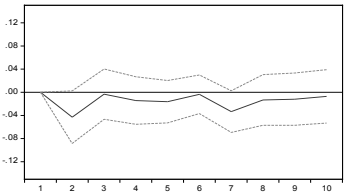


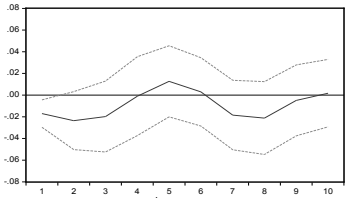
Singapore



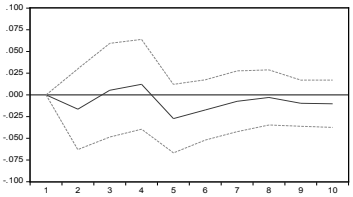


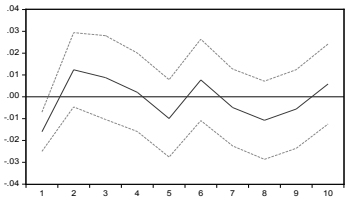
UK





USA





Finally, Table 10 reports the decomposition of the variance of forecast errors for the trade balance and real exchange rates. To gain a better understanding for the sources of long-run variations, the table reports variance decompositions at ten years horizon. The

TABLE 10 - *Decomposition of Variances of Forecast Errors for TB and Q*

	% of *TB* Variance due to *TB*	% of *TB* Variance due to *Q*	% of *Q* Variance due to *Q*	% of *Q* Variance due to *TB*
Australia	62.47	10.94	60.32	31.18
France	41.31	45.83	53.79	28.32
Germany	70.76	9.67	20.90	76.11
Hong Kong	57.31	33.54	70.12	24.58
India	38.22	20.72	29.00	18.62
Italy	45.15	9.57	19.03	47.96
Japan	78.18	13.84	57.68	32.34
Malaysia	34.52	11.28	27.56	10.02
Netherland	75.60	1.49	57.38	23.35
Saudi Arabia	51.55	31.58	55.91	12.42
Singapore	45.67	52.82	92.54	3.98
UK	65.06	10.50	22.72	35.82
USA	54.17	12.89	36.75	36.38

Notes: the values are percentages of variance of 10-step-ahead forecast errors for *TB* and *Q* explained by each other and themselves.

decomposition of the forecast error variance in trade balance due to own shocks and to shocks in exchange rates are reported in columns 2 and 3 respectively[7]. Innovations in trade balance account for a large fraction of the variations in its own forecast errors, ranging from a high of 78.18% for Japan to 34.52% for Malaysia. As for innovations in exchange rates, we can distinguish between three groups: in the low end (less than 20%) are the eight trade partners – Australia, Germany, India, Japan, Malaysia, the Netherlands, U.K. and U.S.A.; the moderate (20% to 40%) includes three partners – Hong Kong, India,

[7] Note that the underlying model includes both domestic and foreign incomes. However, to save space and to concentrate on the main question, we do not report the variance decompositions due to these two variables.

and Saudi Arabia; and the remaining two trade partners – France and Singapore – fall into the high category (more than 40%). Thus, in most cases, innovations in exchange rates play a rather small role in explaining the variability of forecast errors in the trade balance.

The decomposition of forecast error variance for exchange rates due to own innovations and to innovations in trade balance are reported in columns 4 and 5 respectively. Innovations in exchange rates account for a large fraction of variations in their own forecast errors, ranging from a high of 92.54% for Singapore to low of 19.03% for Italy. Innovations in trade balance account for less than 20% of the variations in exchange rates in four cases – India, Malaysia, Saudi Arabia, Singapore; between 20% to 40% in seven cases – Australia, France, Hong Kong, Japan, the Netherlands, U.K. and U.S.A.; and more than 40% in only two cases – Germany, Italy. Thus most long-run variations in the trade balance and exchange rates are explained by innovations in their own underlying processes.

7. Summary and conclusions

The large and growing trade surplus between China and her major trading partners combined with recent global trade imbalances and financial crisis has renewed interest in the long-run relation and short-run dynamics between real exchange rates and the trade balance. A number of critiques have argued that such trade surplus is reflective of government's deliberate policies to undervalue the renminbi, and have recommended major revaluations of the currency to combat the trade surplus. However, such a recommendation implies a stable long-run relation between China's exchange rates and the trade balance. In this study we examine the long-run relation and short-run dynamics between real exchange rate and trade balance using annual bilateral trade data between China and thirteen major trading partners over 1981-2008. We examine the existence of long-run relations using four alternative tests of cointegration: maximum likelihood procedure of Johansen and Juselius to account for endogeneity of variables, Gregory-Hansen's procedure to account for endogenous breaks, Enders and Siklos's momentum threshold autoregressive model to account for non-linearities in the adjustment process, and Pedroni's panel cointegration to account for the small sample size. Irrespective of the empirical method and the choice of trading partner, there is no evidence of significant long-run relation between real exchange rates and the trade balance. We also find no evidence of significant relationship between the two variables in the

short run. Holding domestic and foreign incomes constant, tests of Granger-causality do not find a significant causal relation from real exchange rates to the trade balance, or vice versa. The response of trade balance to innovations in real exchange rates is not significantly different from zero at both short- and long-horizons. Also, most variations of the forecast error in the trade balance are explained primarily by their own innovations. Thus, putting these all together, our results suggest a rather weak relationship between China's trade balance and real exchange rates in both long run and short run. From a policy point of view, the findings suggest that even a major revaluation of renminbi will not necessarily contribute significantly to a reduction in the trade surplus. Thus, recent criticisms of the western countries and their policy recommendations may be unwarranted.

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REFERENCES

Brada, J.C., A.M. Kutan and S. Zhou (1993), "China's Exchange Rate and the Balance of Trade", *Economic Change and Restructuring,* 26(3), 229-242.

Bahmani-Oskooee, M. and T.J. Brooks (1999), "Bilateral J-Curve between U.S. and Her Trading Partners", *Weltwirtschaftliches Archiv,* 135(1), 156-165.

Bahmani-Oskooee, M. and A. Ratha (2004), "The J-Curve: A Literature Review", *Applied Economics,* 36(13), 1377-1398.

Bahmani-Oskooee, M. and Y. Wang (2006), "The J-Curve: China versus Her Trading Partners", *Bulletin of Economic Research,* 58(4), 323-343.

Bahmani-Oskooee, M. and Y. Wang (2008), "The J-Curve: Evidence from Commodity Trade between U.S. and China", *Applied Economics,* 40(21), 2735-2747.

Chan, K.S. (1993), "Consistency and Limiting Distributions of the Least Squares Estimators of a Threshold Autoregressive Model", *Annals of Statistics,* 21(1), 520-533.

Cheung, Y.W. and K.S. Lai (1993), "Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration", *Oxford Bulletin of Economics and Statistics,* 55(3), 313-328.

Enders, W. and P.L. Siklos (2001), "Cointegration and Threshold Adjustment", *Journal of Business and Economic Statistics,* 19(2), 166-176.

Engle, R. and C.W.J. Granger (1987), "Cointegration and Error Correction: Representation, Estimation, and Testing", *Econometrica,* 55(2), 251-276.

Fawley, B. and L. Juvenal (2011), "Why 'Fixing' China's Currency is not a Quick Fix", *The Regional Economist,* 19(2), 4-5.

Garcia-Herrero, A. and T. Koivu (2009), "China's Exchange Rate Policy and Asian Trade", BIS Working Paper No. 282.

Gregory, A.W. and B.E. Hansen (1996), "Residual-Based Tests for Cointegration in Models with Regime Shifts", *Journal of Econometrics,* 70(1), 99-126.

Holtz-Eakin, D. (2003), "The Chinese Exchange Rate and U.S. Manufacturing Employment", Congressional Budget Office Testimony before the Committee on Ways and Means, U.S. House of Representatives.

Johansen, S. and K. Juselius (1990), "Maximum Likelihood Estimation and Inference on Cointegration – with Application to the Demand for Money", *Oxford Bulletin of Economics and Statistics*, 52(2), 169-210.

Johansen, S. (1991), "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", *Econometrica*, 59(6), 1551-1580.

Krueger, A.D. (1993), Exchange Rate Determination, Cambridge University Press: Cambridge.

Lee, J. and M.C. Strazicich (2003), Minimum Lagrange Multiplier Unit Root Test with Two Structural Breaks", *The Review of Economics and Statistics*, 85(4), 1082-1089.

Maddala, G. and S. Wu (1999), "A Comparative Study of Unit Root Tests and a New Simple Test", *Oxford Bulletin of Economics and Statistics*, 61(S1), 631-652.

Narayan, P.K. (2006), "Examining the Relationship between Trade Balance and Exchange Rate: the Case of China's Trade with the USA", *Applied Economics Letters*, 13(8), 507-510.

Pedroni, P. (1999), "Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors", *Oxford Bulletin of Economics and Statistics*, 61(S1), 653-678.

Pedroni, P. (2001), Purchasing Power Parity Tests in Cointegrated Panels", *Review of Economics and Statistics*, 83(4), 727-731.

Pedroni, P. (2004), "Panel Cointegration; Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis", *Econometric Theory*, 20(3), 597-625.

Toda, H.Y. and T. Yamamoto (1995), "Statistical Inferences in Vector Autoregressions with Possibly Integrated Processes", *Journal of Econometrics*, 66(1-2), 225-250.

Wane, A., S. Gilbert and S. Dibooglu (2004), "Critical Values of the Empirical F-Distribution for Threshold Autoregressive and Momentum Threshold Autoregressive Models", Department of Economics Discussion Papers, Southern Illinois University Carbondale.

Wang, Z. (2007), "Real Effect of Exchange Rate in RMB and Trade Balance Surplus in China", *Canadian Social Science*, 3(5), 91-96.

Wei, W. (1999), "An Empirical Study of the Foreign Trade Balance in China", *Applied Economics Letters*, 6(8), 485-490.

Yaya, M.E. and X. Lu (2010), "The Short-Run Relationship between Real Effective Exchange Rate and Balance of Trade in China", SSRN Working Paper Series.

Zhang, Z. (1998), "Does Devaluation Improve China's Balance of Trade?", *Economia Internazionale/International Economics*, 51(3), 437-445.

Zhang, Z. (1999a), "China's Exchange Rate Reform and Its Impact on the Balance of Trade and Domestic Inflation", *Asia Pacific Journal of Economics & Business*, 3(2), 4-22.

Zhang, Z. (1999b), "Foreign Exchange Rate Reform, the Balance of Trade and Economic Growth: An Empirical Analysis for China", *Journal of Economic Development*, 24(2), 143-162.

Zhang, Z. and K. Sato (2009), "How Effective is the Renminbi Devaluation on China's Trade Balance", Working Paper Series, The International Centre for the Study of East Asian Development: Kitakyushu.

Zivot, E. and D.W.K. Andrews (1992), "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis", *Journal of Business & Economic Statistics*, 10(3), 251-270.

ABSTRACT

We investigate long-run relations and short-run dynamics between China's bilateral trade balance and real exchange rates with thirteen major trading partners over 1981-2008. Maximum likelihood tests of cointegration reveal no evidence of significant long-run relationship between the two variables. These findings are robust to endogenous breaks in long-run relations, non-linear relations in the adjustment process, and the sample size. There is also no evidence that real exchange rates have a significant effect on the trade balance in the short run. These are reflected in tests of Granger-causality, impulse response functions and variance decompositions. Thus, there is no support for the view that China's large and growing trade surplus is a reflection of undervalued RMB. As such, policies aimed at revaluing RMB do not necessarily contribute to a significant decline in China's trade surplus.

Keywords: Asymmetric Adjustments, Causality, Cointegration, Real Exchange Rates, Structural Breaks, Trade Balance, Unit Roots

JEL Classification: F31, F41

RIASSUNTO

La bilancia commerciale e i tassi di cambio: ulteriori evidenze dalle bilance commerciali bilaterali cinesi

In questo studio vengono esaminate la relazione di lungo periodo e le dinamiche di breve periodo tra le bilance commerciali bilaterali cinesi e i tassi reali di cambio con i maggiori tredici partner nel periodo 1981-2008. I test di cointegrazione *maximum likelihood* non evidenziano significative relazioni di lungo periodo tra le due variabili. Questi risultati resistono a rotture endogene nelle relazioni di lungo periodo, in quelle non lineari nel processo di aggiustamento e alla dimensione del campione. Inoltre non emerge evidenza che i tassi reali di cambio abbiano un effetto significativo sulla bilancia commerciale nel breve periodo. Ciò è riscontrato nel test di causalità di Granger, nelle funzioni di risposta agli impulsi e nelle scomposizioni di varianza. Non c'è quindi sostegno alla tesi per cui il grande e crescente surplus commerciale cinese sarebbe una conseguenza della sottovalutazione del renminbi. Pertanto politiche volte alla sua rivalutazione non porterebbero necessariamente ad una significativa riduzione del surplus commerciale cinese.