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Does dollar-pegging matter? A closer look at US trade deficits with China and Germany

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ABSTRACT
China and Germany are comparable in terms of having persistent trade surplus with the USA, but they differ in how their currencies are valued. By invoking the China–Germany comparison, this paper finds that there is weak, if any, statistical association between the US trade deficit and the exchange rate. This finding is robust to long-run vs. short-run horizon, without vs. with an instrumental variable, and in-sample fitting vs. out-of-sample forecasting. This paper predicts that the US trade deficits with China and Germany will continue to rise in the presence of a recovery in the US economy.

KEYWORDS China; exchange rate; trade; cointegration

JEL CLASSIFICATION C32, F31, F32, O53

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1. Introduction

The contribution of the slow-moving exchange rate of Chinese currency yuan (renminbi) to the US trade deficit with China is a hot debate topic. Yuan's *de facto* crawling peg to US dollar is interpreted by some as the evidence that China is a currency manipulator[1]. This paper attempts to add new insight to that debate by posing the question of what would happen to the US trade deficit if China let its currency float freely. This counterfactual is unlikely to occur in the near future, but this what-if exercise is crucial in uncovering the causality.

We provide an answer by contrasting the US trade deficit with China to the deficit with Germany. There are two rationales for this identification strategy. First, the US has run sustained and substantial trade deficits with both countries in recent years[2]. Second, these two countries differ in the regime of exchange rate. We argue that the value of the German currency, the Deutsche mark before 1 January 2002, and the euro since then, is largely driven by market forces[3]. China is different. The Chinese central bank sets a daily reference midpoint rate and currently allows yuan to trade 2% above or below that target[4]. The Chinese central bank maintains this trading band through constant buying and selling of foreign exchange reserves.

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In terms of experimental design, let pegging to the dollar or actively intervening in the foreign exchange market be the treatment. China is, therefore, in the treatment group and Germany is in the control group of no active intervention. Our goal is to deduce the effect of the treatment on the US trade deficit. The null hypothesis is that the treatment has no effect, and it cannot be rejected if the marginal effect of the exchange rate on the US trade deficit with China is not significantly different from Germany.

Toward this end, we estimate multivariate seemingly unrelated regressions (SUR) over rolling windows, in which the log real trade deficit is regressed on the log real exchange rate. The SUR enables us to test the equal-exchange-rate elasticities of US trade deficit with China and Germany. The rolling windows can account for potential structural breaks and avoid arbitrarily splitting the whole sample into before-treatment and after-treatment sub-samples. In particular, we are interested in those post-2002 windows in which we see more variation in the Chinese exchange rate.

Besides our identification strategy, there are three econometric issues to address. First, most of the time series that we use are nonstationary, and we must consequently guard against spurious regressions via the test of cointegration. Second, the *t*-statistic in the cointegration regression used by Engle and Granger (1987) follows a nonstandard distribution. Instead, we apply the dynamic ordinary least squares (OLS) estimator of Stock and Watson (1993) to which traditional inference can apply . Finally, there may exist reverse causation from the trade deficit to exchange rate[5]. In order to remove the simultaneity bias, we provide an instrumental variable (IV) estimate, where the IV is the Chinese housing market index. We assume the housing index affects the exchange rate through the financial account[6], but has no direct effect on the trade deficit in goods. Some of our specifications involve the lagged values of exchange rates, and this approach can also alleviate the simultaneity issue.

We investigate how the exchange rate affects the trade deficit in both the long run and the short run. In terms of econometrics, we consider both Engle–Granger type cointegration regression, and the error correction model (ECM); the former uses the *level* of the exchange rate as the regressor, whereas the latter utilizes the *difference* of the exchange rate. By comparing the two models, we can verify the conjecture that what matters is not the appreciation of Chinese currency, but how fast the appreciation is.

Finally, we evaluate the out-of-sample predictive power of the exchange rate for US trade deficit. The baseline model includes both the exchange rate and US real GDP. In one alternative model, the exchange rate is dropped, while in the other model US real GDP is omitted. We examine how omitting those two variables affects the forecasting error, in an attempt to extract signal from noise.

Our work ties into a growing literature on Chinese exchange rate and US trade imbalance. For example, our results are in line with Zhang and Wan (2007) and Zhang and Wan (2008), which employ structural vector autoregression (VAR) analysis and find that changes in the exchange rate bear little on the trade balance. Based on the purchasing power parity model and the shadow price of foreign exchange model, Chou and Shih (1998) provide evidence that the Chinese Government has adopted an exchange rate policy that maintains official exchange rates close to equilibrium levels, so the exchange rate is not the primary determinant for the trade imbalance with USA. Similarly, Zhang (2001) shows that the economic reforms in China have brought the real exchange rate closer to the equilibrium level. By contrast, Bahmani-Oskooee and Ardalani (2006) apply the error-correction modeling to the data at the industry level, and show that the real yuan–dollar exchange rate has played a significant role in the trade imbalance. Xu (2008)

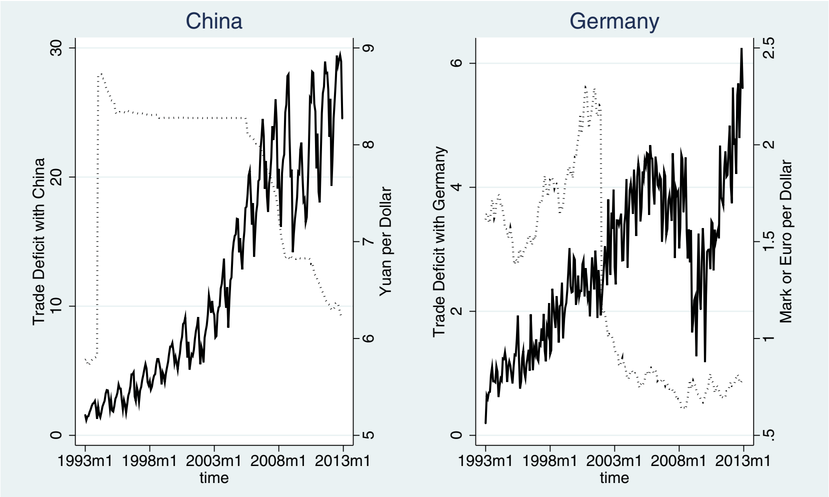

Figure 1. US trade deficits (solid line) and exchange rates (dotted line).

detects a statistically significant long-run relationship between the yuan–dollar exchange rate and the US trade deficit, even though in short run the link between them is not evident.

Some authors investigate the US trade imbalance more generally, not just with China. For example, Bahmani-Oskooee and Wang (2007) examine monthly import and export data from 66 industries in the the USA, and show that in the long run real depreciation of the dollar can improve export earnings of many US industries, but have no significant effect on most importing industries. Chinn and Ito (2007) investigate the medium-term factors that drive the current account imbalance using a model that controls for factors related to institutional development. The authors find the link between the developed equity markets and current account deficits.

This paper distinguishes itself by invoking the China–Germany contrast, the IV estimation, and the comparison of out-of-sample forecasting errors. Moreover, including the recent Great Recession in the analysis can facilitate the identification by adding the variability of key variables.

2. Data

We obtain nominal US trade in goods[7] with China and Germany (in billions of US dollars, not seasonally adjusted) from the US Census Bureau[8]. The Chinese housing index (Constant Quality Price Index for Newly-Built Private Housing in 35 Major Chinese Cities) is provided by the authors of Wu, Gyourko, and Deng (2010). All other series are downloaded from Federal Reserve Economic Data[9]. Except for US real GDP, US federal budget deficit, and Chinese housing index, we have monthly observations from January 1993 to December 2012 with a sample size of 240[10]. We obtain the monthly US real GDP and monthly US federal budget deficit by dividing the quarterly data by three [11].

The solid line in the left panel of Figure 1 represents the nominal US trade deficit in goods with China[12]. Overall, the series is above zero with an upward trend. In December 2012, the trade deficit was 24 billion dollars. The noticeable up-and-down fluctuation within a year suggests seasonality – examining data carefully reveals that imports from

China typically peak before the Thanksgiving holiday. Later, we will include quarterly dummy variables in regressions to account for this seasonality[13].

The dotted line denotes the nominal exchange rate quoted as yuan per dollar (on the right vertical axis). We see an upward jump from about 6 to 8 in the early 1990s[14], which translates to a substantial devaluation of the yuan or appreciation of the dollar. The yuan's rate remained largely fixed until 2005, and since then there has been a gradual revaluation of yuan or depreciation of dollar. In December 2012, one dollar traded for 6.2 yuan.

The right panel of Figure 1 plots the US trade deficit with Germany and the nominal rate of German currency. The USA–Germany trade deficit shows a similar pattern as the USA–China deficit but with a smaller magnitude; it is this similarity that motivates our identification strategy. There was a downward shift in the German rate in January 2002 caused by the euro's introduction. We will, therefore, run the Germany regressions separately, before and after 2002.

If exchange rates play the dominant role for trade, then we expect to see the exchange rate and trade deficit move in the *same* direction. That is, other factors being secondary, a depreciating US dollar will result in more competitive American goods and a smaller US trade deficit. Figure 1, however, shows only short-lived evidence for that projection: the co-movement of the exchange rate and trade deficit is visible only between 1995 and 2002 for Germany. The two series actually moved in the *opposite* direction after 2005 for China and between 2002 and 2008 for Germany. After 2008, the euro's exchange rate seems to wander independently of the US–Germany trade deficit.

Taken literally, Figure 1 is indicative of little, if any, association between the nominal exchange rate and trade deficit. US GDP, though, stands out as a prominent factor. Notice that the trend of rising trade deficits was temporarily halted in early 2000s (for Germany, in particular) and was reversed between 2008 and 2010. Both periods saw recessions in USA, during which shrinking income could afford fewer imports. The 2008–2009 downturn was more severe and was associated with a more visible improvement in the US trade balance.

Table 1 reports some descriptive statistics. Several facts are noteworthy. On average, the USA–China trade deficit is about four times as large as the USA–Germany deficit. The Chinese yuan has appreciated by more than 30% in nominal terms from its trough. Finally, the augmented Dickey–Fuller (ADF) test with 12 lags and a linear trend indicates that all series except Chinese CPI are non-stationary. Given this result, we should be wary of spurious regressions and other econometric difficulties.

3. Empirical methodology

3.1. Seemingly unrelated regressions

Our identification entails comparing the following two seemingly unrelated regressions (SUR)

$$\mathrm{lrtdc}_t = \beta_0^c + \beta_1^c lrec_t + \beta_2^c \mathrm{control} + e_{ct} \tag{1}$$

$$\mathrm{lrtdg}_t = \beta_0^g + \beta_1^g lreg_t + \beta_2^g \mathrm{control} + e_{gt} \tag{2}$$

where the dependent variables are (natural) log real US trade deficit in goods with China (lrtdc) and Germany (lrtdg). The key regressors are log real exchange rates of Chinese yuan (lrec) and German currency (lreg). Table 2 provides detailed information about

Table 1. Descriptive statistics.

Variables	Statistics					
	N	Mean	SD	Min	Max	ADF(12)
Trade deficit with China (tdc)	240	12.2	8.6	1.2	29.4	− 2.3
Trade deficit with Germany (tdg)	240	2.8	1.3	0.2	6.2	− 1.9
Nominal yuan exchange rate (ec)	240	7.7	0.8	5.7	8.7	0.2
Nominal euro/mark exchange rate (eg)	240	1.2	0.5	0.6	2.3	− 2.1
China CPI (cpic)	240	84.9	13.1	47.5	109.3	− 4.1
Germany CPI (cpig)	240	89.9	7.9	75.8	105.0	− 2.2
US CPI (cpiu)	240	185.0	26.5	142.8	231.6	− 1.9
US M2 (m2u)	240	5952.1	2019.3	3406.8	10400.1	0.0
US federal budget (fbu)	240	− 31.6	49.7	− 159.7	70.4	− 2.1
US real GDP (rgdpu)	240	4302.1	622.6	3141.1	5144.6	− 1.5
China housing index (mchi)	108	2.2	0.9	0.9	3.7	− 2.8

Note:
a. *N* is the sample size; mean is the sample mean; SD is the standard deviation; Min is the minimum; Max is the maximum. ADF(12) is the augmented Dickey–Fuller unit root *t* test with 12 lags and a linear trend. The corresponding critical value at the 5% level is −3.41.
b. The sample spans from January 1993 to December 2012 for all variables except mchi.

Table 2. Explanations of variables: January 1993–December 2012.

Variable	Explanation
Original variables	
tdc	US trade deficit with China (billions of dollars)
tdg	US trade deficit with Germany (billions of dollars)
ui	US import from China (billions of dollars)
ec	Nominal yuan exchange rate: yuan per dollar
eg	Nominal euro/mark exchange rate: euro/mark per dollar
cpic	Consumer price index in China
cpig	Consumer price index in Germany
cpiu	Consumer price index in USA
m2u	M2 money stock in USA (billions of dollars)
fbu	Federal government budget in USA (billions of dollars)
rgdpu	Real gross domestic product in USA (billions of chained 2009 dollars)
Regressands	
lrtdc	Log real US trade deficit with China = log(tdc) - log(cpiu)
lrtdg	Log real US trade deficit with Germany = log(tdg) - log(cpiu)
lrui	Log real US import from China = log(ui) - log(cpiu)
Regressors	
lrec	Log real yuan exchange rate = log(ec) + log(cpiu)-log(cpic)
lreg	Log real euro/mark exchange rate = log(eg) + log(cpiu)-log(cpig)
rfbu	Real federal budget in USA = fbu/cpiu
lm2u	Log M2 money stock in USA = log(m2u)
lrgdpu	Log real gross domestic product in USA = log(rgdpu)
lrgdpg	Log real gross domestic product in Germany
dqi	Dummy variable; =1 for the *i*-th quarter of a year
Instrument	
mchi	China housing market index

Note: Physical euro coins and banknotes entered into circulation on 1 January 2002. So eg is Deutsche mark per dollar before 1 January 2002, and euro per dollar after. This also applies to lreg.

constructing those variables. For instance, the log real exchange rate of Chinese yuan is computed as

$$\text{lrec} = \log(\text{nominal exchange rate yuan per dollar}) + \log(\text{US CPI}) - \log(\text{China CPI}) \tag{3}$$

After lrec declines, Chinese goods become more expensive relative to American goods or the yuan appreciates in real terms. We use the real exchange rate instead of the nominal rate for three reasons. First, what matters for trade is the real rate since it embodies the price differential. Second, because the CPI and wage tend to move together, using the real exchange rate can to a large extent account for the difference in labor costs, of which accurate data are hard to obtain in China. Finally, using the real rate can take into account 'pass-through', i.e. the change in the nominal rate may lead to disproportional change in the real rate if the prices of traded goods are set in dollars.

Notice that (1) and (2) involves the log–log specification. Hence, β_1^c and β_1^g measure elasticities and are readily comparable. We are especially interested in two null hypotheses:

$$H_0^{\text{weak}} : \beta_1^c = \beta_1^g \tag{4}$$
$$H_0^{\text{strong}} : \beta_1^c = \beta_1^g = 0 \tag{5}$$

The elasticities are the same under H_0^{weak}. If we assume the German exchange rate is free-floating, then H_0^{weak} implies that the yuan's rate effectively behaves like a floating one, or alternatively that the yuan's pegging to dollar (sluggish movement) has no effect. Testing H_0^{weak} enables us to statistically answer that counterfactual question of what if China had adopted the floating rate, so it is the key to our identification strategy. Under the stronger H_0^{strong} the exchange rates have a statistically irrelevant impact on the trade deficit.

We refer to economic theory and stylized facts when selecting control variables. For example, a simple open-economy model discussed in Chapter 6 of Mankiw (2013) suggests log real US GDP (lrgdpu) and real US federal government budget deficit (rfbu)[15] – the former is suggested by Figure 1 to be potentially important, and the latter can be used to verify the relationship of the trade deficit and the budget deficit (twin deficits)[16]. The theory of the J-curve[17] necessitates the inclusion of lagged values of exchange rates. Ho, Zhang, and Zhou (2014) investigate the effect of quantitative easing on financial markets. This paper extends that finding by examining how log US M2 money stock (lm2u) affects the trade deficit. The neutrality of money holds if lm2u is insignificant. Because the R-squared is 0.99 when regressing lm2u onto a linear trend, lm2u can also approximate other trending unobserved factors. Finally, we include quarterly dummy variables to capture the seasonality.

3.2. Dynamic GLS estimates using rolling windows

One may argue that there may be other structural changes in addition to the euro's introduction, such as China's joining WTO in 2001 and the revaluation of the yuan in 2005. Searching and accounting for an unknown number of breaks can be technically challenging and is not the focus of this study. We instead choose to run sequential regressions using rolling windows. More explicitly, we first run regressions using the first window consisting of $t = 1, 2, \ldots,$ windowsize observations. We then move to the second window $t = 2, 3, \ldots,$ windowsize $+ 1$, and run the regressions again. We continue until the last window reaches the end of the sample. If using the dummy variables is the parametric approach of accounting for breaks, then running rolling-window regressions can be seen as the more flexible non-parametric approach. The window size is chosen so that the window is long enough for the cointegration test to have adequate power but not too long to mute structural breaks.

We have as yet ignored the fact that the t-value in the Engle–Granger type cointegration regression does not follow the normal distribution even asymptotically. The tests of our hypotheses (4) and (5) likewise follow nonstandard distributions. In order to correctly apply the conventional statistical inference, we need to utilize the dynamic OLS estimator proposed by Stock and Watson (1993), which in our multivariate SUR setting becomes the dynamic GLS estimator. Basically, we need to include the leads and lags of first-differenced regressors. See Stock and Watson (1993) for details.

3.3. *Instrumental variable estimation*

In the cointegration regression, the super-consistency of the OLS estimator holds even when the regressor and the error term are correlated[18]. Although in large sample we can downplay the endogeneity issue, in small sample it might still help to resolve the endogeneity by calling for instrumental variables. For our problem, the endogeneity may arise due to, say, the reverse causation from the trade deficit to the exchange rate – a rising trade surplus with the USA will put the Chinese Government under greater political pressure to revaluate yuan.

Any IV should satisfy two requirements: being strongly correlated with the endogenous regressor, and being excluded from the structural model. The IV that we choose is the Constant Quality Price Index for Newly-Built Private Housing in 35 Major Chinese Cities, or housing index for short. Ho, Zhang, and Zhou (2014) document that, with soaring housing prices, speculative funds or 'hot money' may rush into China[19]. All else equal, those capital inflows are likely to prompt an appreciation of the yuan. This means the first requirement (relevance) is met. Meanwhile, we assume the housing index affects the US trade deficit only *indirectly* through its effect on the exchange rate. Statistically, this is equivalent to saying that the housing index can be excluded from the structural regression of the trade deficit once the exchange rate has been controlled for. This assumption is plausible in that most new-housing-related goods such as cement, steel, and furniture are produced and consumed within China. Put differently, the housing index affects the exchange rate only through the financial account, not the current account[20].

3.4. *Error correction model*

By using the level of the exchange rate as a regressor, we have so far only examined its long-run effect on the trade deficit. A related question is, what if the *difference* of exchange rate is the regressor? Equivalently, what if it is the growth rate of the exchange rate that matters? After all, the recurring complaint of US Government officials is not that Chinese currency has not appreciated, but rather that the appreciation has been too slow. From the standpoint of econometrics, one may question whether the previous results are driven by the fact that both the trade deficit and US GDP are trending and wonder whether the same results carry over to the de-trended series.

To address both concerns, we turn to the ECM given by

$$\Delta \text{lrtdc}_t = \alpha_0^c + \gamma^c \text{ecc}_{t-1} + \alpha_1^c \Delta \text{lrec}_t + \alpha_2^c \Delta \text{lrgdpu}_t + \alpha_3^c \Delta \text{control} + u_{ct} \quad (6)$$

$$\Delta \text{lrtdg}_t = \alpha_0^g + \gamma^g \text{ecg}_{t-1} + \alpha_1^g \Delta \text{lreg}_t + \alpha_2^g \Delta \text{lrgdpu}_t + \alpha_3^g \Delta \text{control} + u_{gt} \quad (7)$$

where Δ is the difference operator. Basically, we remove the low-frequency component or trend by taking differences, which is appropriate for a difference-stationary process[21]. The error correction terms are ecc and ecg, obtained as the residuals of the Engle–Granger cointegration regressions. Cointegration exists if $\gamma \neq 0$, otherwise the ECM reduces to a regression in differences. Here we choose the unrestricted approach by allowing $\gamma \neq 0$.

Notice that α_1 and α_2 measure the effects of the change of exchange rate and US GDP, and the t-values of α_1 and α_2 follow standard normal distributions, see Sims, Stock, and Watson (1990). In this new setting, the counterparts of (4), (5), (14), and (15) are

$$H_0^{\text{weak}} : \alpha_1^c = \alpha_1^g \quad (8)$$

$$H_0^{\text{strong}} : \alpha_1^c = \alpha_1^g = 0 \quad (9)$$

$$H_0^{\text{weak}} : \alpha_2^c = \alpha_2^g \quad (10)$$

$$H_0^{\text{strong}} : \alpha_2^c = \alpha_2^g = 0 \quad (11)$$

3.5. Out-of-sample forecasting

Alternatively, the role of the exchange rate can be evaluated by its predictive power for the trade deficit – the exchange rate is important if excluding it leads to substantial increase in the forecasting error. We compare three models. The benchmark or 'long' model is a seemingly unrelated ADL with both log real exchange rate and log US real GDP used as regressors:

$$\text{lrtdj}_t = \phi_{j,0} + \phi_{j,1} \text{lrej}_t + \phi_{j,2} \text{lrgdpu}_t + \phi_{j,3} \text{control}_t + \eta_{j,t}, \quad (Mo1) \quad (12)$$

where j is the index for countries, $j = c, g$, and the control variables include the lagged regressand, log US M2 stock, US real federal budget deficit, and quarterly dummies. The long model (12) nests two 'short' models as special cases: the second model *Mo*2 keeps US GDP (lrgdpu) but drops the exchange rate (lre*j*), and the third model *Mo*3 keeps the exchange rate but drops US GDP.

The three models are fitted using the rolling windows. More explicitly, suppose the m-th window ends at the t_m-th observation. We fit the models using that window and then compute the squared one-step-ahead forecasting error as

$$\text{squared forecastint error}_{t_m+1} = (\text{lrtdj}_{t_m+1} - \widehat{\text{lrtdj}}_{t_m+1})^2 \quad (13)$$

where $\widehat{\text{lrtdj}}_{t_m+1}$ denotes the out-of-sample predicted value. Note that the $(t_m + 1)$-th observation is not used for the in-sample fitting, so it is the (pseudo) out-of-sample forecasting we are after.

3.6. Reduced form vector autoregression (VAR)

We also estimate separately for China and Germany a three-variable reduced form VAR[22] that includes (deseasonalized) real trade deficit, real US GDP and real exchange rate, with and without the exogenous US M2 money stock and US budget deficit. All variables are in

Table 3. ADL estimates.

	China(lrtdc)				Germany(lrtdg)	
	1993 m1–2012 m12				Before 2002	After 2002
	(1)	(2)	(3)	(4)	(5)	(6)
lrec	0.092	0.032	0.223	0.063		
	(0.160)	(0.120)	(0.349)	(0.101)		
rfbu	−0.025	−0.421***	−0.375***	−0.235***	−0.286	0.309*
	(0.076)	(0.094)	(0.099)	(0.083)	(0.346)	(0.178)
lm2u	0.322	−0.305	−0.154	−0.210	−1.434*	−0.132
	(0.205)	(0.195)	(0.215)	(0.168)	(0.810)	(0.342)
lrgdpu	4.020***	5.071***	4.773***	2.349***	5.016***	1.378
	(0.401)	(0.366)	(0.407)	(0.425)	(1.152)	(1.098)
lrec_{t-1}			−0.848*			
			(0.475)			
lrec_{t-2}			0.026			
			(0.473)			
lrec_{t-3}			0.518			
			(0.476)			
lrec_{t-4}			0.215			
			(0.348)			
lrtdc_{t-1}				0.565***		
				(0.058)		
lreg					−0.002	0.373
					(0.281)	(0.252)
lrtdg_{t-1}					0.101	0.521***
					(0.083)	(0.075)
Quarterly dummies	No	Yes	Yes	Yes	Yes	Yes
N	240	240	236	239	107	132
AIC	−127	−268	−264	−347	−67	−53
Engle–Granger test	−7.01	−9.78	−10.08	−14.51	−7.11	−6.40

Note: Each column represents a separate regression. The standard error are in parentheses. $***$, $**$, and $*$ denote significance at 1%, 5% and 10% levels, respectively.

difference in light of the nonstationarity. The lag length is chosen based on the Schwarz's Bayesian information criterion (SBIC). In particular, we are interested in testing the null hypothesis that the real US GDP or real exchange rate does not Granger cause the US trade deficit.

4. Results

We first estimate (1) and (2) separately. Table 3 reports the estimated coefficients for each model. For China, we use the whole sample and start with a static model without quarterly dummy variables (reported under column (1)). We then add quarterly dummies (column (2)) and four lagged values of the log real exchange rate (column (3)). Column (3) is the distributed lag (DL) model. The DL model's benefit is that it allows for the lingering effect of exchange rates on the trade deficit. The disadvantage is that, in the presence of persistent lingering, the DL model must include many lagged values, resulting in undesirable multicollinearity and consequent imprecision.

As a parsimonious alterative, the autoregressive distributed lag (ADL) model is reported in column (4); this model includes only the first lag of the dependent variable lrtdc_{t-1}. This ADL model is algebraically equivalent to a DL model with infinite number of lagged regressors. In addition to simplicity, there are two more rationales for using the lagged regressand: it can serve as a proxy for those time-invariant or slow-evolving unobserved factors[23], and it can also substantially reduce the serial correlation left in the error term.

According to the Akaike information criterion (AIC), the ADL model outperforms the static model, which outperforms the DL model[24]. To save space, for Germany we only report the ADL model: one specification uses the observations before 1 January 2002 (column (5)), and the other uses observations after that date (column (6)). This sample-splitting is due to Germany's switching from the mark to the euro.

For all regressions, we report the Engle–Granger test of no cointegration, and the spurious regression can be ruled out in favor of cointegration in all cases. The existence of cointegration has two econometric implications. First, the coefficient estimate is superconsistent, and the *t*-value follows a nonstandard distribution. Second, the serial correlation in the error term becomes inconsequential[25], even when the lagged regressand is present.

Based on the conventional statistical inference, Table 3 illustrates that the exchange rate (lrec or lreg) is statistically insignificant in all specifications. By contrast, US real GDP (lrgdpu) is statistically significant in five out of six cases. Consider column (4). In that regression, lrec is neither economically nor statistically significant, but lrgdpu is significant at the 1% level. Furthermore, the trade deficit is shown to be elastic with respect to US total income, in that one percentage rise in real US GDP leads to more than 2% increase in trade deficit. This finding is consistent with Figure 1 and accords well with the discussion of Mankiw (2013).

One may be bothered by the fact that the coefficient of lrec in column (3) 0.223 differs markedly from column (4) 0.063. This contrast is merely superficial; the former coefficient is the artifact of multicollinearity, as evidenced by the fact that the standard error (reported in parentheses) of lrec in the DL model is over triple that of the ADL model.

The US M2 money stock is shown to be insignificant in most cases. That means 'hot money', if it really exists, mainly affects the financial markets rather than the traded-goods sector. There is some evidence that the US Government's budget may affect the trade deficit with China and pre-euro Germany, as smaller budget deficits are associated with smaller trade deficits.

Table 4 reports the results when the ADL versions of regressions (1) and (2) are jointly estimated, using the separated subsamples before and after 2002. Thanks to unidentical regressors, in theory there is efficiency gain associated with the generalized least squares (GLS) estimates relative to the equation-by-equation OLS estimates. Despite the potential gain in precision, the statistical significance of lrec, lreg and lrgdpu remains unchanged. The efficiency gain can be seen by comparing the standard error of lrgdpu: it falls from 1.098 in column (6) of Table 3 to 1.055 in column (10) of Table 4.

There is also direct benefit of joint estimation – now we can formally test the hypotheses (4) and (5). Neither hypothesis can be rejected at the 5% level. Thus, the key message of Tables 3 and 4 is that exchange rates, whether free-floating or slow-crawls, are statistically irrelevant at least for the USA–China and USA–Germany deficits. The 'elephant in the room' instead turns out to be the US real GDP.

By using a windowsize of 96 (eight years) and the subsamples before and after 2002, Figures 2 and 3 plot the series of Stock–Watson's dynamic *t*-values, which asymptotically follow the standard normal distribution. The *t*-value of the log real exchange rate is

Table 4. SUR-ADL estimates.

	Before 2002		After 2002	
	China	Germany	China	Germany
	(7)	(8)	(9)	(10)
lrec	0.148		0.390	
	(0.183)		(0.260)	
$lrtdc_{t-1}$	0.481***		0.362***	
	(0.093)		(0.077)	
lreg		0.008		0.395
		(0.267)		(0.240)
$lrtdg_{t-1}$		0.116		0.528***
		(0.079)		(0.072)
rfbu	−0.281	−0.286	−0.230***	0.300*
	(0.253)	(0.331)	(0.076)	(0.172)
lm2u	−1.186***	−1.426*	−0.119	−0.141
	(0.451)	(0.772)	(0.265)	(0.330)
lrgdpu	3.870***	4.940***	4.008***	1.420
	(0.836)	(1.099)	(0.643)	(1.055)
Quarterly dummies	Yes	Yes	Yes	Yes
N	107	107	132	132
Engle–Granger test	−3.62	−7.24	−5.17	−6.33
P-value of testing (4)	.667		.987	
P-value of testing (5)	.719		.084	

Note: Each column represents a separate regression. The standard error are in parentheses. ***, **, and * denote significance at 1%, 5% and 10% levels, respectively.

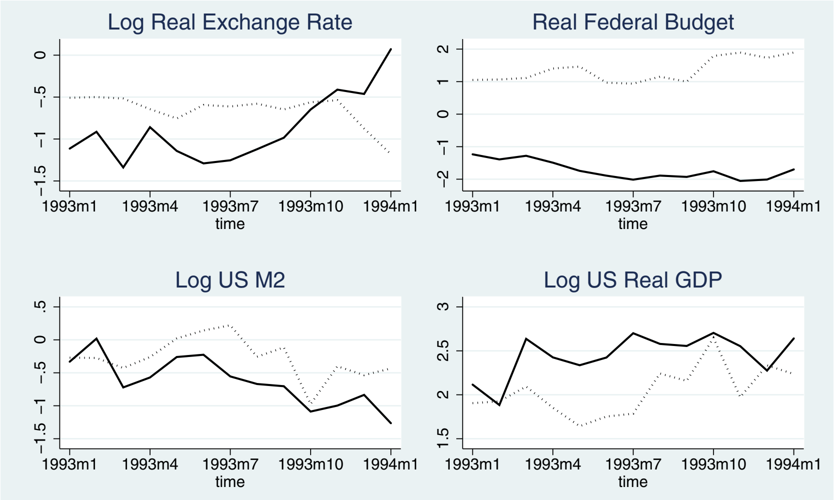

Figure 2. *T*-values of coefficients (before 2002, window = 96). China (solid line) and Germany (dotted line).

always inside the (−1.96, 1.96) interval for China (solid line), and is outside that interval only immediately after 2002 for Germany (dotted line). By contrast, the *t*-value of log US real GDP moves outside the (−1.96, 1.96) interval much more frequently. For China after 2002, the log US real GDP is always significant at the 5% level.

In Figure 4, we see slightly different patterns with the window size being 60 (five years): the shorter window makes the *t*-value series less smooth, and for Germany the log real

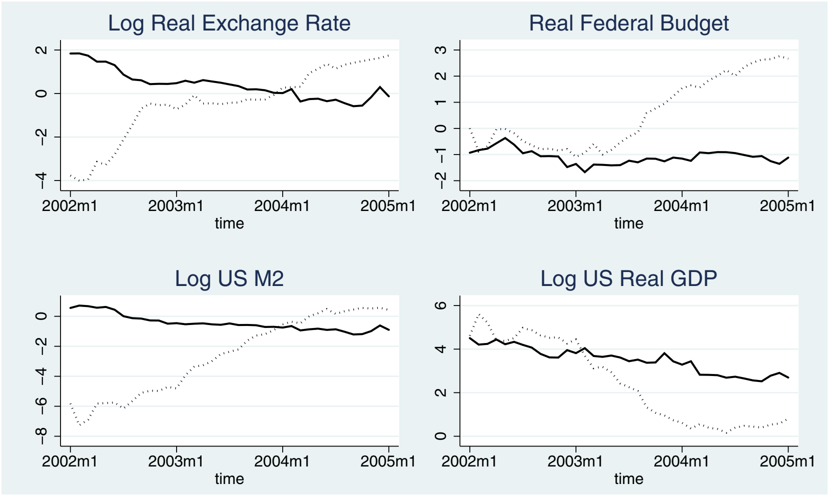

Figure 3. *T*-values of coefficients (after 2002, window = 96). China (solid line) and Germany (dotted line).

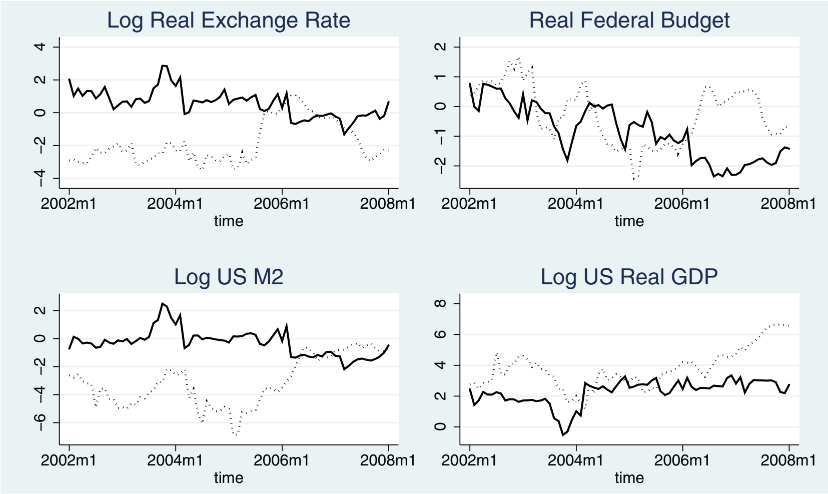

Figure 4. *T*-values of coefficients (after 2002, window = 60). China (solid line) and Germany (dotted line).

exchange rate appears to be significant with higher frequencies. The pattern becomes clearer in Figure 5, which plots the series of p-values of testing hypotheses (4) and (5) based on the dynamic GLS estimates. A dash horizontal line that corresponds to the p-value of 0.05 is drawn. As expected, it is easier to reject the stronger hypothesis (5) than (4). The two hypotheses cannot be rejected in almost all windows after 2006. For completeness, Figure 5 also plots the p-value of testing the following two hypotheses:

$$H_0^{\text{weak}} : \beta_{21}^c = \beta_{21}^g \tag{14}$$

$$H_0^{\text{strong}} : \beta_{21}^c = \beta_{21}^g = 0 \tag{15}$$

where β_{21}^c and β_{21}^g are the coefficients of log real US GDP in (1) and (2), respectively. Hypothesis (14) states that the income elasticity of the trade deficit is the same across

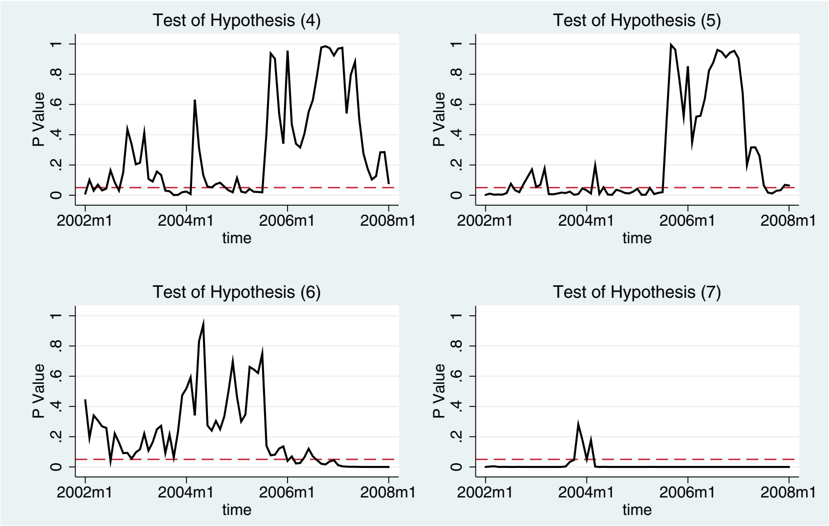

Figure 5. After 2002, window = 60.

Table 5. IV estimates.

	IV1 (11)	IV2 (12)
lrec	0.108	0.519
	(2.680)	(0.520)
$lrtdc_{t-1}$	0.419***	0.412***
	(0.142)	(0.118)
rfbu	−0.215	−0.225
	(0.256)	(0.206)
lm2u	−0.339	
	(1.831)	
lrgdpu	3.847***	3.706**
	(1.109)	(1.614)
N	108	108
First-stage *F* test	1.742	21.068

Note: Each column represents a separate regression. The standard error are in parentheses. ∗∗∗, ∗∗, and ∗ denote significance at 1%, 5% and 10% levels, respectively.

China and Germany, and it cannot be rejected most of the time according to Figure 5. Hypothesis (15) implies a zero income elasticity, and it is overwhelmingly rejected except in several windows.

In short, Figure 5 reaffirms the previous finding that US real GDP is the dominant factor for the US trade deficit.

The IV estimates of the two ADL specifications are reported in Table 5. With log US M2 money stock included in the regression, the IV is weak (the first-stage *F*-test is less than 2), but it becomes a strong IV after dropping lm2u from the first-stage and second-stage regressions. This change results from the collinearity between lm2u and the housing index[26]. Regardless of whether the IV is weak or strong, the estimated effect of log real

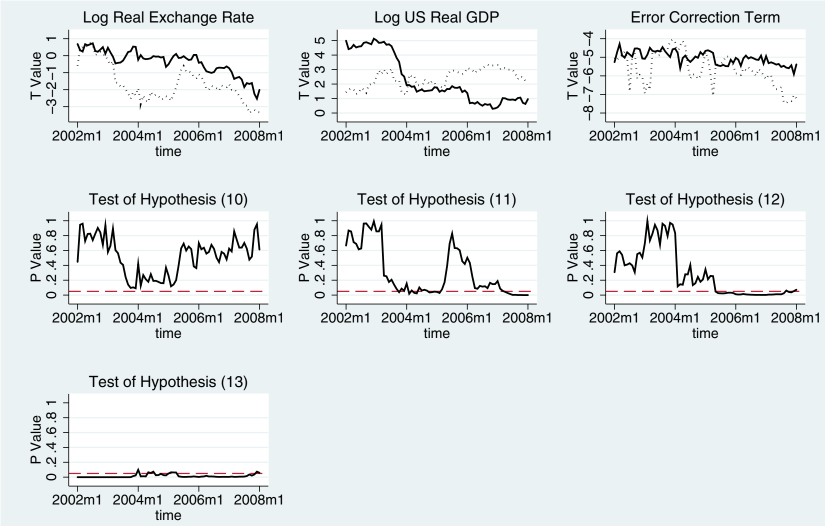

Figure 6. ECM (after 2002, window = 60).

exchange rate on the trade deficit is insignificant, whereas the log US real GDP is significant. Moreover, the economic significance of the log US real GDP in Table 5 is close to that in column (9) of Table 4[27].

The similarity between the OLS and IV estimates reaffirms that the endogeneity issue is minor in the cointegration setting. Alternatively, it may also imply the exogeneity of exchange rate (insignificant reverse causation), which is in line with the possibility that the Chinese government does not take seriously the US Government's appeal for a stronger yuan.

We estimate (6) and (7) jointly, with rolling windows of size 60 applied to the post-2002 subsample. Figure 6 plots the t-values of γ, α_1 and α_2, and the p-values of testing (8),(9),(10), and (11).

Figure 6 shows no qualitative change relative to Figures 4 and 5. The t-value of log real exchange rate is most often inside the $(-1.96, 1.96)$ band, while the t-value of log US real GDP is outside that band in at least half of the windows. Only hypothesis (11) can be overwhelmingly rejected.

Figure 6 does provide some new insight. Notice that the t-value of the error correction term in the Germany regression (7) (denoted by dotted line) tends to be more significant[28] than that in the China regression (6) (denoted by solid line). That means the error correction term plays a more important role for the USA–Germany deficit. In other words, the tendency to revert to the long run equilibrium of the USA–Germany deficit is stronger than the USA–China deficit.

Table 6 reports the mean and standard deviation of the forecasting error. One striking finding is that dropping the exchange rate actually *improves* the forecasting: the mean of squared forecasting error (MSFE) decreases from 0.0095 of *Mo*1 to 0.0086 of *Mo*2 for China, and from 0.0471 to 0.0451 for Germany. Meanwhile, the standard deviation of forecasting error decreases for China, implying that the forecast without the exchange rate is tighter.

Table 6. Summary of squared out-of-sample forecasting error.

	China			Germany		
	Mo1	Mo2	Mo3	Mo1	Mo2	Mo3
Mean	0.0095	0.0086	0.0112	0.0471	0.0451	0.0660
Standard deviation	0.0186	0.0141	0.0223	0.0737	0.0737	0.1041
P-value of DM test Mo2 = Mo1		0.5370			0.0432	
P-value of DM test Mo3 = Mo1		0.4510			0.2728	

Table 7. *P*-value of Granger causality test.

Regressand	Excluded	China		Germany	
US trade deficit	US real GDP	0.00	0.00	0.09	0.23
US trade deficit	Real exchange rate	0.16	0.19	0.18	0.20
US real GDP	US trade deficit	0.03	0.58	0.45	0.58
US real GDP	Real exchange rate	0.68	0.30	0.71	0.41
Real exchange rate	US trade deficit	0.02	0.01	0.45	0.53
Real exchange rate	US real GDP	0.11	0.09	0.55	0.54
Exogenous Control		No	Yes	No	Yes

Note: The reported number is the *p*-value for testing the null hypothesis that the excluded variable does not Granger cause the regressand in a three-variable reduced form VAR, with and without exogenous control variables of US M2 money stock and budget deficit.

Unlike the exchange rate, ignoring the US GDP worsens the forecast: the MSFE rises from 0.0095 of *Mo*1 to 0.0112 of *Mo*3 for China, and from 0.0471 to 0.0660 for Germany. In this case, there is increase in the standard deviation of forecasting error.

Table 6 also reports the test of equal predictive accuracy proposed by Diebold and Mariano (1995) (DM test), where six lagged values are used in calculating the long-run variance for the Bartlett kernel. The null hypothesis of equal predictive accuracy can be rejected at the 5% level only when *Mo*2 is compared to *Mo*1 for Germany. The significance of the DM test remains unchanged when four lagged values are used.

Overall, these findings indicate that, once again, the dominant factor for US trade deficit is US GDP. The value-added of the forecasting analysis is that we can now distinguish the signal and noise – the signal is informative for the future but the noise is not. For the US trade deficit, the US GDP is very likely to be the signal and the exchange rate the noise.

Table 7 reports the *p*-value of Granger causality test in the three-variable reduced form VAR, with and without exogenous control variables of US M2 money stock and budget deficit[29]. We can reject the null hypothesis that the US real GDP does not Granger cause the US trade deficit at the 10% level for three out of four cases, but can reject the null hypothesis that real exchange rate does not Granger cause the US trade deficit in none of cases. Again, this finding indicates that it is the US real GDP rather than the real exchange rate that has in-sample explanatory power for the US trade deficit.

Moreover, there is evidence that the US trade deficit Granger causes the US–China real exchange rate, reflecting that China is under the pressure to adjust the exchange rate in the presence of accumulating trade imbalance. Finally, it is shown that the US–China exchange rate can also be affected by the US real GDP.

Table 8. Robustness check I.

	China(2005 m7–2012 m12)				Germany(lrtdg)	
	$Y = lrtdc$		$Y = lrui$		Before 2002	After 2002
	(13)	(14)	(15)	(16)	(17)	(18)
lrec	−2.402	0.596	−1.383	0.449		
	(1.499)	(0.408)	(1.150)	(0.303)		
rfbu	−0.410***	−0.291**	−0.333***	−0.215**	−0.293	0.281
	(0.123)	(0.116)	(0.094)	(0.087)	(0.347)	(0.183)
lm2u	−0.388	−0.060	−0.046	0.109	−1.250	−0.236
	(0.436)	(0.308)	(0.334)	(0.228)	(0.847)	(0.373)
lrgdpu	6.987***	4.223***	6.232***	3.459***	5.558***	1.072
	(1.092)	(1.130)	(0.838)	(0.911)	(1.355)	(1.182)
lrec_{t-1}	3.442		1.458			
	(2.438)		(1.870)			
lrec_{t-2}	0.062		0.495			
	(2.427)		(1.862)			
lrec_{t-3}	2.030		1.644			
	(2.399)		(1.840)			
lrec_{t-4}	−2.667*		−1.884			
	(1.597)		(1.225)			
lrtdc_{t-1}		0.476***				
		(0.106)				
lrui_{t-1}				0.509***		
				(0.104)		
lreg					−0.015	0.339
					(0.282)	(0.257)
lrtdg_{t-1}					0.101	0.520***
					(0.083)	(0.075)
lrgdpg					−1.643	0.708
					(2.153)	(1.000)
Quarterly dummies	yes	yes	yes	yes	yes	yes
N	90	90	90	90	107	132
AIC	−162.73	−180.66	−210.45	−233.68	−66.42	−52.12
Engle–Granger test	−4.14	−2.98	−3.15	−2.47	−6.84	−6.16

Note: Each column represents a separate regression. The standard error are in parentheses. ∗∗∗, ∗∗, and ∗ denote significance at 1%, 5% and 10% levels, respectively.

5. Robustness check

Another questionable margin is whether it makes sense to run a regression when the Chinese exchange rate is strictly pegged to dollar, which if literally implemented implies no variation in the nominal exchange rate. Another concern is that we have so far ignored the Chinese and German GDP. In this section, we provide several robustness checks to allay such concerns. First, columns (13) and (14) in Table 8 report the DL and ADL estimations for China but only using observations after July 2005, a sub-sample in which China abandoned the strict peg and the variation of the yuan's exchange rate is substantial. Chinese GDP presumably only affects US export to China, so columns (15) and (16) report the DL and ADL regressions in which the log real US imports from China (lrui), not the trade deficit, is used as the dependent variable[30]. Finally, columns (17) and (18) extend columns (5) and (6) in Table 3 by including log real German GDP (lrgdpg).

Comparing Table 8 to Table 3 reveals few qualitative differences. The insignificance of the exchange rate remains, as does the significance of US GDP. Furthermore, German GDP is shown to be insignificant for US trade deficit. Given that the standard deviations

Table 9. Robustness check II.

	China (1993 m1–2012 m12), $Y = lrtdc$					
	(19)	(20)	(21)	(22)	(23)	(24)
$lrtdc_{t-1}$	0.538***	0.543***	0.531***			
	(0.060)	(0.060)	(0.061)			
$lrtdc_{t-3}$				0.197***	0.192***	0.180***
				(0.064)	(0.063)	(0.065)
lrgdpu	2.678***	2.544***	2.707***	4.052***	4.037***	4.177***
	(0.484)	(0.504)	(0.538)	(0.565)	(0.570)	(0.609)
d	0.614	0.558	0.833	−0.228	0.187	0.583
	(0.876)	(0.890)	(0.958)	(1.006)	(1.016)	(1.090)
lrec	0.173			0.079		
	(0.157)			(0.178)		
$d \times lrec$	−0.191			0.114		
	(0.302)			(0.347)		
$lrec_{t-3}$		0.185			0.208	
		(0.156)			(0.179)	
$d \times lrec_{t-3}$		−0.172			−0.029	
		(0.307)			(0.350)	
$lrec_{t-6}$			0.221			0.294
			(0.157)			(0.180)
$d \times lrec_{t-6}$			−0.266			−0.165
			(0.329)			(0.374)
Controls	Yes	Yes	Yes	Yes	Yes	Yes
AIC	−347.90	−348.67	−343.31	−283.87	−285.38	−283.14

Note: Each column represents a separate regression. The standard error are in parentheses. ∗∗∗, ∗∗, and ∗ denote significance at 1%, 5% and 10% levels, respectively.

are 2780 and 11,203 for US export and import with China, respectively, we think that the role of Chinese GDP is likely to be marginal as well.

We next focus on China and investigate the structural break of China's abandoning the strict peg in July 2005. Table 9 generalizes Table 3 in three ways. First, a dummy variable d is defined for that break date, and interaction terms of d and (lagged) log real exchange rates are generated. The idea is to allow for different intercepts and slopes before and after the break date. Second, the third and sixth lag values of log real exchange rates are used to account for more lagged response of trade deficit to the exchange rate. Finally, the third lag of the dependent variable is considered for the same reason.

Table 9 indicates that the first lag regressand has more explanatory power than the third lag. The US real GDP is significant while the dummy variable d is insignificant in all specifications. Most importantly, allowing for time-varying intercepts and slopes does not affect the role of log real exchange rate. That variable and its interactions with the structural-break dummy are insignificant in all specifications.

The last finding is consistent with the view that, even though China can manipulate its nominal exchange rate, it cannot do the same for its real exchange rate, and it is the real rate that matters for the trade deficit. Before July 2005, the standard deviation of log Chinese real exchange rate is 0.068. In the same period, the standard deviation of the difference of log CPIs in USA and China is 0.099. So during the strict-pegging period, the majority of the variation in the real exchange rate comes from the inflation differential. After all, a booming export sector had led to rising Chinese wages, which caused the real appreciation of Chinese currency despite the nominal peg to dollar. After July 2005, the

Table 10. Robustness check III.

	China(lrtdc)				Germany(lrtdg)	
	1993 m1–2012 m12				Before 2002	After 2002
	(1a)	(2a)	(3a)	(4a)	(5a)	(6a)
lrec	0.128	0.058	0.319	0.113		
	(0.162)	(0.123)	(0.356)	(0.104)		
rfbu	0.001	−0.379***	−0.320***	−0.162*	−0.204	0.365**
	(0.076)	(0.096)	(0.099)	(0.084)	(0.341)	(0.174)
lm2u	0.405*	−0.230	−0.045	−0.045	−1.397*	0.025
	(0.208)	(0.201)	(0.217)	(0.171)	(0.828)	(0.330)
lrgdpu_{t-1}	3.836***	4.919***	4.551***	1.880***	4.920***	0.814
	(0.407)	(0.376)	(0.409)	(0.448)	(1.174)	(1.051)
lrec_{t-1}			−0.963**			
			(0.484)			
lrec_{t-2}			−0.011			
			(0.482)			
lrec_{t-3}			0.731			
			(0.484)			
lrec_{t-4}			0.088			
			(0.356)			
lrtdc_{t-1}				0.598***		
				(0.062)		
lreg					−0.036	0.295
					(0.281)	(0.250)
lrtdg_{t-1}					0.103	0.532***
					(0.084)	(0.075)
Quarterly dummies	No	Yes	Yes	Yes	Yes	Yes
N	239	239	236	239	107	132
AIC	−121	−256	−255	−335	−66	−53
Engle–Granger test	−6.93	−10.06	−10.26	−15.42	−7.23	−6.50

Note: Each column represents a separate regression. The standard error are in parentheses. $***$, $**$, and $*$ denote significance at 1%, 5% and 10% levels, respectively.

standard deviations became 0.112 and 0.028, respectively. So in the peg-crawl period, the variation in the real exchange rate is mostly driven by the adjustment of nominal rate. Regardless of the main source of real appreciation of yuan, Table 9 reinforces the conclusion that the exchange rate plays only a secondary role in explaining the USA–China trade deficit.

Finally, Table 10 duplicates Table 3 using the US real GDP in the previous year as the regressor other than the contemporaneous one, in order to mitigate the reverse causality from the trade deficit to the real GDP in the same year. The main findings remain qualitatively unchanged: the real exchange rate is shown to be statistically insignificant in all specifications, while the US real GDP is significant in five out of six specifications. Overall, this robustness implies that the issue of reverse causality from the trade deficit to US GDP may be minor, which is consistent with the fact that the US trade deficit with China only counts about 2% of US GDP in recent years.

6. Conclusion and discussion

Our introduction posited the counterfactual question of what would happen to the US trade deficit if China allowed the yuan to float. The answer is that, as long as the US economy reverts back to its pre-crisis level, the trade deficit with China would not decrease.

In other words, the Chinese currency's slow movement does not matter as much as rising US GDP for the US trade deficit. This conclusion is drawn from the cointegration analysis, the IV estimation, and the out-of-sample forecasting.

The finding that the US–China trade deficit is insensitive to the exchange rate is actually not surprising, given the two countries' positions in the global value chain – China has the comparative advantage in labor intensive goods and the USA in capital intensive goods. In short, the sustained USA–China trade deficit is largely due to the fact that Chinese goods complement rather than compete with American goods.

We can also reconcile our findings with the Triffin Dilemma which implies that the perpetual US trade deficit has something to do with the fact that US dollar is a reserve currency. In other words, the world financial system would run out of liquidity and become frozen if the USA ran balanced trade.

What is surprising is the marginal role of the exchange rate in the USA–Germany trade deficit. This finding in some sense supports a recent view that the euro is undervalued from the perspective of the USA–Germany trade balance and that Germany has been benefiting from the cheap euro. Recently, there is an article titled 'Germany's trade surplus is a problem' on Ben Bernanke's blog[31]. The main idea of that paper is that although the euro can be at the equilibrium level for all 19 euro-zone countries as a group, given the wage and production cost in Germany, the euro–dollar exchange rate can be too weak, leading to substantial Germany trade surplus. Put differently, the comparatively weak euro serves as unexpected benefit to Germany that comes with the participation in the currency union. Had Germany still used the Deutsche mark, its currency would probably be much stronger than the present euro, which could reduce the cost advantage of German exports.

This provides an alternative interpretation of our results, namely, in terms of causing trade imbalance, the undervalued euro behaves effectively the same as the crawling peg Chinese currency. In another article titled 'Lean on me,' the author suggests that a faster growth rate of wage in Germany may partially solve the trade imbalance caused by the undervalued euro[32].

There are two possible ways to extend this study. One is to find two countries with persistent trade *deficits* with the USA, but one country uses the fixed exchange rate or actively intervenes the market, and the other does not. Alternatively, we may look for countries that switch from one exchange rate regime to another and examine the impact on the respective trade balance. While the USA–China trade deficit does not permit the use of these alternative strategies, its likely prominence in the fast approaching 2016 US presidential election warrants this opening approach. Other researchers may be able to complement our work from these angles.

Notes

1. For example, Mitt Romney, the Republican's 2012 presidential nominee, said 'I've watched year in and year out as companies have shut down and people have lost their jobs because China has not played by the same rules, in part by holding down artificially the value of their currency.' See *New York Times* report titled 'A Tightrope on China's Currency' on October 22, 2012.
2. We do not consider Japan even though it also has persistent trade surplus with the US. In contrast to the Euro, the Japanese currency is a sovereign currency. It is therefore more likely to be subject to intervention by the Japanese Government.
3. Before 2002, German Government's interventions in the foreign exchange market may have occasionally occurred, but, if so, they were not typical. Since 2002, there has been little evidence that

the European central bank has systematically intervened in the foreign-exchange market just for Germany's sake.

4. In December 2001, China joined the World Trade Organization with the promise to gradually adjust its currency level. In July 2005, the yuan was revalued by 2.1 %, which can be seen as the beginning of the so-called 'managed floating exchange rate based on market supply and demand with reference to a basket of currencies.'
5. USA may put greater pressure on China to revalue the yuan if its trade deficit is rising quickly. For instance, see the *Bloomberg* article on 14 Nov 2011 titled 'Obama Says "Enough is Enough" on China Currency Valuation'.
6. Speculative funds or 'hot money' may flow to China in the expectation of rising house values, see Ho, Zhang, and Zhou (2014). Everything else equal, these capital inflows will lead to appreciation of yuan.
7. We do not use the trade data in service for two reasons: first, they are less accurate than goods data and second, they are largely irrelevant if the focus is on 'manufacturing jobs'.
8. http://www.census.gov/foreign-trade/balance/
9. http://research.stlouisfed.org/fred2/
10. The starting and ending dates are chosen to maximize the data availability. For instance, the Chinese CPI reported in FRED starts in January 1993.
11. We want to thank an anonymous referee who suggests using the Denton method of obtaining monthly data, see Denton (1971) for details. We find no qualitative difference in our results.
12. There is no qualitative change in the pattern if the ratio of trade deficit to US GDP is put on the vertical axis.
13. Alternatively, one may apply the Census Bureau X-12-ARIMA method to remove the seasonal components, see Makridakis, Wheelwright, and Hyndman (1998) for details.
14. In 1994, China ended the dual exchange rate system, and the official rate was merged with the market-determined rate in the swap center.
15. We do not take log because rfbu is typically negative but can be positive.
16. The federal budget deficit does not necessarily cause the trade deficit. For instance, the change in the private saving will offset the change in the public saving under Ricardian equivalence. Total saving would then remain unchanged, as would the trade deficit.
17. See page 674 of Feenstra and Taylor (2014) for discussion.
18. Consider a simple regression $y_t = \beta x_t + e_t$, $(t = 1, 2, \ldots, T)$, where y_t and x_t are integrated of order one. We can show that the probability limit of the OLS estimate is $\hat{\beta} = \beta + \frac{T^{-2}\sum x_t e_t}{T^{-2}\sum x_t^2} = \beta + o_p(1)$, which holds even if the regressor is endogenous $T^{-1}\Sigma x_t e_t \neq 0$. Notice that for a non-stationary regressor x_t we need to multiply $\sum x_t^2$ by T^{-2} to achieve non-divergence. See Phillips and Durlauf (1986) for more discussion.
19. Despite the official capital control of Chinese Government, some capital inflow and outflow may occur through underground channels, see Martin and Morrison (2008) for instance.
20. Consider a hypothetical transaction. Suppose an American hedge fund manager decides to invest 1 million dollars to buy Chinese real estate – he imports a Chinese asset. He then gives the Chinese seller a check drawn on a US bank, an export of US asset to China. According to the double-entry bookkeeping principle, in the US. Balance of Payments both entries fall within the financial account (with different signs), and the current account is unaffected. This example is modified from the discussion on page 592 of Feenstra and Taylor (2014).
21. We do not consider things like Hodrick–Prescott filter because our interest is not on the business cycle.
22. Conducting a full-scale structural VAR analysis is beyond the scope of this paper, see Zhang and Wan (2007) and Zhang and Wan (2008) for a structural VAR analysis of a similar topic. Reduced form VAR has the advantage of assuming away (possibly incorrect) identification restrictions and disadvantage of failing to provide orthogonal error terms, so we do not report the impulse-response function.
23. An example is the Chinese labor cost, for which we do not have data.
24. Using Bayesian information criterion (BIC) generates the same ranking.
25. See Phillips and Durlauf (1986) for detailed discussion.
26. The *R*-squared is 0.84 when regressing lm2u onto mchi.
27. The IV regression only uses 108 observations because the housing index starts in 2004.

28. Rigorously speaking, the *t*-value of γ follows a nonstandard distribution under H_0: $\gamma = 0$. Here, we ignore that complexity and still use the conventional critical values.
29. We thank an anonymous referee for suggesting the VAR analysis.
30. We hesitate to include Chinese GDP in the regression because many believe the official data have been inflated, see Bosworth and Collins (2008) for instance.
31. http://www.brookings.edu/blogs/ben-bernanke/posts/2015/04/03-germany-trade-surplus-problem
32. http://www.economist.com/news/finance-and-economics/21695913-europes-weak-economic-recovery-worryingly-dependent-exports-lean-me

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