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Renewed estimation of a single equation for the Chinese Renminbi

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Abstract

This paper provides evidence on the consistency of the determination of the Chinese real effective exchange rate (REER) over time. Especially, we validate cointegration between the REER and a set of fundamentals using recent developments in model selection. Error correction model (ECM) path dependence in model selection is addressed by using the General-To-Specific (GETS) approach enabling us to obtain empirically constant and encompassing ECM. As inference in finite samples is commonly of concern, statistics' distributional properties for cointegration tests are estimated by Monte Carlo simulations. The final specification of the model is compatible with the natural real exchange rate of Stein (1994). We briefly study the implications of our findings in terms of foreign exchange policy.

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Keywords: Exchange Rate, Equilibrium Value, GETS, Global imbalances

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1 Introduction

In recent years, China has exhibited impressive external accounts and growth rates but has operated few adjustments in the value of its currency. This situation has given ground to fierce opposition to the country's foreign exchange policy. In this paper, we contribute to the debate by estimating a Behavioural Equilibrium Exchange Rates (BEER) which is a single equation of the Real Effective Exchange Rate (REER) and the fundamentals. This type of BEER regressions have been conducted by many authors since the seminal work of Clark and MacDonald (1997 and 2000). However, in the case of emerging markets, the econometric estimation of a reduced-form equation for a single currency often stumbles on the size of the sample available. Especially, results might be too sensitive to specific observations. Also the estimation might not be empirically constant. An alternative to single currency estimation is panel estimation which enables researchers to work on large samples. However, it is well known that panel estimations erase country specific information, as mentioned by the CGER (IMF 2006) and Dunaway, Leight & Li (2009), producing, often, per se, large deviations from equilibrium values. The aim of this paper is to introduce a small-sample-proof estimation of the BEER of a single currency using very recent developments in cointegration techniques. This type of technique should become the norm, especially for emerging markets, which suffer the most from the sample size problem. In this paper, and as noted by Clark and MacDonald (2000), we show that with appropriate econometric methods and proper specification, results become empirically constant, data-coherent and robust to the most recent and harshest diagnostic tests. In line with the literature, we apply co-integration techniques to estimate the BEER equation for the Chinese currency. Doing so, we introduce two innovations. The first innovation lies in the estimation of the error correction model. It is well known that many alternative single-equation models might be obtained depending upon the model search path followed. In this paper, ECM path dependence in model selection is addressed by using the General-To-Specific approach (GeTS) developed by Hendry and Krolzig (2001). This is a multi-path model selection algorithm. It relies on diagnostic tests of congruence to simplify models in a multi-step process. The final model is of course parsimonious but it is also encompassing as among all possible models it is the one that presents the best set of responses to diagnostic tests. Particularly, the final model is empirically constant. The second innovation lies in the final test of the co-integration hypothesis. The distribution of ECM statistics is commonly of concern in small samples, particularly in the tails. This might produce false signals, biasing the test toward the acceptance of co-integration. We correct this bias using the Ericsson and MacKinnon (2002) methodology for finite sample inference about co-integration in which statistics' distributional properties are estimated by Monte Carlo simulations. Finally, specifying the deterministic component appropriately enables us to enlarge the usual sample used in the literature from 1980 to 2008. We do this testing for several dates of optimal trend break using methodologies developed by Perron (1989), Perron and Zhu (2005) and Zivot and Andrews (1992). Especially, we find that the devaluation of the renminbi, in 1994, is not the optimal time location of the structural break in the series.

Final results validate the existence of a co-integration relationship between the REER of China and a set of fundamentals. Especially, using the harshest diagnostic tests, we confirm that, since the early 1980s, the behaviour of the currency is intimately related to such fundamentals as productivity, absorption and the terms of trade. This specification is in line with the natural exchange rate of Stein (1994) and shares factors with many theoretical models. Capitalizing on this result we show that increasing efforts of internal consumption might justify a significant revaluation of the currency. Also we are able to quantify the shock on absorption necessary to create domestic circumstances as safe as to weather such an adjustment.
As the econometric methodology we use is recent and technical we detail it extensively in the body of this article. As BEER is not an innovative model, its conclusions are well known. Therefore we only briefly comment on the loadings of the BEER equation and their consequences in terms of political economy. This paper is organized in four sections. Following this introduction, in section two, we characterize the theoretical framework of this work by going further into details in the BEER methodology. Also we list and compare the most significant works available on Chinese data. In section three, we develop the quantitative analysis of the data introducing innovations in the estimation of the ECM and the final test of co-integration. Also we analyze the results and propose a brief economic interpretation. We conclude in a fourth section.

2 Reduced-form equation of equilibrium exchange rates

As noted by Isnard (2007), the decade that followed the breakdown of the Bretton Woods System in 1971 gave rise to "a 'heroic age' of exchange rate theory" and to many econometric estimates of reduced form exchange rate equations. Innovative models have been proposed by Mussa (1984), Edwards (1989), Elbadawi (1994), Stein (1994) and Faruqee (1995). The Behavioural Equilibrium Exchange Rate (BEER) developed by MacDonald (1997) and Clark and MacDonald (2000) encapsulates many of these models. That is because it is a purely statistical methodology. Often it serves as a general framework to estimate single equation exchange rate equilibriums. Egert (2004) made an elegant presentation of the BEER as follow:

The theoretical underpinning of the BEER approach rests on the uncovered real interest rate parity (URIP):

$$\mathbb{E}_t(q_{t+1}) - q_t = r_t - r_t^* \tag{1}$$

where r_t, and r_t^* represent the domestic and foreign ex ante real interest rates, $E_t(q_{t+1})$ stands for the expected real exchange rate in t for $t+1$, and q_t is the observed real exchange rate. It is a function of the expected value of the real exchange rate in t for $t+1$ and the ex ante real interest differential.

$$q_t = E_t(q_{t+1}) - (r_t - r_t^*) \tag{2}$$

$E_t(q_{t+1})$ can be assumed to be the outcome of the expected values of the fundamentals, so that

$$q_t = E_t(\bar{x}_{t+1}) - (r_t - r_t^*) \tag{3}$$

where $\bar{x}$ is the vector of fundamentals. In practical terms, the real exchange rate can be written as the function of long and medium-term (x) fundamentals and short-term variables (z):

$$q_t = q_t(\bar{x}_t, \bar{z}_t) \tag{4}$$

Building on this general equation, approaches to estimate reduced-form exchange rate equations differ in their identification of fundamentals and the methodology used to generate their long-term values. A general theoretical framework to determine fundamentals is given by Elbadawi (1994) who capitalizes on Edwards (1989). Starting off with an identity for nominal domestic absorption, Elbadawi (1994) derives equations for demand and supply for non-traded goods enabling him to state an equilibrium in the non-traded goods sector. Linking domestic prices to international prices, he generates an equation for the real exchange rate that ensures instantaneous equilibrium in the non-traded goods market for given levels of some exogenous and policy fundamentals[1]:

$$q_t = q_t\left(\frac{A}{Y}(+), TOT(?), Tx(+), Tm(+), \frac{EXPgn}{EXPg}(+), \frac{EXPg}{Y}(?)\right) \tag{5}$$

Elbadawi (1994) notes that the above solution suggests that higher and sustainable levels of the domestic absorption ratio $\frac{A}{Y}$, foreign trade taxes on exports Tx and imports Tm and the ratio of government expenditure on non-tradable goods to government expenditure on tradable goods $\frac{EXPgn}{EXPg}$ are consistent with equilibrium real exchange rate appreciation (sign + next to the variables). Meanwhile, the effects due to the terms of trade and the total government expenditure ratio could not be signed a priori. However, consistent empirical regularity shows that improving terms of trade and higher government expenditure tend to lead to RER appreciation because the income effect of the terms of trade improvement usually dominates the substitution effect, and governments tend to have a higher propensity to spend on non-traded goods than the private sector[2]. To estimate the model, we use the following linearized version of the equation:

$$Log(q_t) = \alpha + \beta 1 log(TOT) + \beta 2 log(OPEN) + \beta 3 log(A/GDP)$$
$$+\beta 4 log(\frac{curr.G.EXP}{G.EXP}) + \beta 5 log(\frac{G.EXP}{GDP}) \quad (6)$$

As government expenditure in non-tradable goods is difficult to assess in the case of China, we measure the Balassa Samuelson effect using productivity differential as is common in the literature (see table 1). The equation (6) can be reduced to:

$$Log(q_t) = \alpha + \beta 1(TOT) + \beta 2(OPEN) + \beta 3(Tcons) + \beta 4(Prod) \quad (7)$$

Where:

1. Log(q_t) is the logarithm of the Real Effective Exchange Rate. This is the CPI-based version calculated by the IMF and available in the World Economic Outlook database. A positive change reflects an appreciation for the home country's currency.

2. ToT is the logarithm of the terms of trade measured by the ratio of export prices to import prices for China. Data are from the World Development Indicators database of the World Bank. Following Elbadawi (1994), we do not have particular expectations concerning the sign of this variable.

3. OPEN is a measure of the degree of openness of the economy. It is usual to proxy OPEN by the ratio of the sum of exports and imports to GDP as in Egert (2004). Also, one can find imports only to GDP ratio as a proxy of OPEN as in Krueger, Kamar and Carlotti (2009). We have tested both variables using the IMF International Financial Statistics database. The expected sign of OPEN is mixed in the literature as changes in the degree of openness might be related to larger imports which might depreciate the REER, larger exports which might appreciate the REER or larger imports and exports which might have no consequence for the REER.

4. Tcons stands for total consumption. It is the logarithm of the ratio of final consumption expenditure to GDP, A/Y in equation (5). The relationship between the REER and absorption is likely to be positive. Tcons describes the internal equilibrium in the economy between investments, consumption and savings. As a consequence it might be approximated by the resulting external position measured by net foreign assets as in Funk & Rahn (2005) and Bénassy-Quéré et al. (2006). However, in the case of China which maintains capital controls, NFA appears largely endogenous to exchange rates. Also, the interest rates differential,

which appears sometimes in single equations, suffers from the simultaneity bias due to capital controls in China.

5. Prod is a measure of productivity to account for the Balassa Samuelson effect. The relative ratio of consumer price index to producer price index is a usual proxy for productivity in the literature (see table 1). It estimates relative prices between non-tradable and tradable goods. However, in the case of China, the existence of numerous government controlled prices may introduce a bias in this measure. In this paper, we use Real GDP to employment ratio as an alternative measure of productivity. Data are from the OECD database. It is largely admitted in the literature that rising productivity tends to appreciate the REER[3].

Even though China might be considered as a small economy during most of the period under study, it is difficult to state any prior belief on the statistical exogeneity of the variables. Actually, the importance of planification in China calls rather for a lack of exogeneity and the presence of simultaneity bias. This has consequences in term of inference and forecast and must be tested.

Before presenting the results of our research, we summarize the most recent estimations of a single equation for the Chinese Renminbi in table 1. Especially we report the dataset used for the estimation.

Table 1: Behavioural Equilibrium Exchange Rates

The Behavioral Equilibrium Exchange Rate (BEER) as defined by Clark and McDonald (2000) can be considered as a statistical approach. It does not ensure equilibrium but verifies the existence of a relationship in the long run between the REER and a set of fundamentals. Actually, all econometric estimates of a single equation of the REER and its fundamentals can be considered encapsulated in the BEER methodology.

	Variable	Sample	Exogenous variables	(-)/(+)valuation
MacDonald & Dias (2007)	REER	G3 and 10 emerging economies 1988 Q1 to 2006 Q1 (Q)	Net exports ((-) for G3 and (+) for full sample); Real Interest Rate differential (-); GDP Per Capita differential (+); Terms of Trade differential (+)	-8% to -42% (2006 Q1)
Stolper & Fuentes (2007)	REER	12 Asian countries	Productivity differentials (+); Terms of Trade differentials (+)	-4.8% (2007)
Bénassy-Quéré & al. (2006)	REER	15 G20 countries and World 1981 Q1 to 2004 Q3 (Q)	Productivity differential*((-) for 15 countries and (+) for World); NFA position(-)	-31% (15 countries) to -45% (World) (2004 Q3)
Funk & Rahn (2005)	REER	China 1994 Q1 to 2002 Q4 (Q)**	Productivity differential(+)*; NFA position (+)	-3% (2002 Q2)
Wang (2004)	REER	China 1980 to 2003 (Y)	Productivity Change; NFA; Openess (Signs not stated explicitly)	-5% (2003)
Stolper & Fuentes (2007)	USD/CNY	12 Asian countries	Productivity differentials (+); Terms of Trade differentials (+)	-5% (2007)
Bénassy-Quéré & al. (2006)	USD/CNY	15 G20 countries and world 1981 Q1 to 2004 Q3 (Q)	Relative productivity*; NFA position (Signs not stated explicitly)	-30% to -59% (2004)
Funk & Rahn (2005)	USD/CNY	China 1994 Q1 to 2002 Q4 (Q)**	Productivity differential* (+); NFA position** (+)	-12% (2002)
Coudert & Couharde (2007)	USD/CNY	21 emerging countries 1980 Q1 to 2004 Q4 (Q)	Real exchange rate***(+)	-18% (2002)

For each study in column one, we report the type of exchange rate, REER or bilateral against the USD, in column two. In column three we report the dimension of the panels in the cross section and time series. We also report the frequency of observation: Q for quarterly and Y for yearly data. In column four we report the significant exogenous variables with the sign of the loadings. Finally we report the estimated disequilibrium (under (-) or over (+) valuation) and the year of assessment in column five.*Relative productivity of tradable goods versus non-tradables goods. This ratio is approximated by the ratio of consumer price index to producer price index.**Funke and Rahn (2005) started to estimate their model in a sample covering the period 1985 to 2002 but reduced it to 1994 to 2002 due to the presence of a break in the series in 1994 following the devaluation of the currency.***Authors defined the real exchange rate as the nominal exchange rate (indirect quote) multiplied by the ratio of the final demand price index in the emerging country / final demand price index in the US.

In this paper, we re-estimate a well known model, as in equation 7, using recent cointegration techniques. We use a sample starting in Q1 1980 enabling us to estimate the equation on 116 quarterly observations for China only. Since 1980, China has experienced several changes of foreign exchange policy, especially a devaluation in Q1 1994. These changes have generated breaks in the series of observations that we must account for. We do this implementing the methodologies for times series with known structural break as in Perron (1989), Perron and Zhu (2005) and unknown structural breaks as in Zivot and Andrews (1992). When necessary, yearly data are transformed to quarterly data using the quadratic optimization procedure [4]. Doing so, we clearly impose a model on the data-generating process of the missing observations. The risk is to generate artificial autocorrelation in the series and thus to obtain estimators with non minimum variance. Therefore, we must be particularly vigilant when reading diagnostic tests. However, for highly persistent fundamentals as productivity or absorption this should not be too constraining. Also the problem may be addressed correcting the variance covariance matrix for autocorrelation and heteroskedasticity as proposed by Newey and West (1987)[5]. As a consequence:

- we significantly extend the usual sample under study from the early 1980s to 2008 covering the period of large reserve accumulation that revived debate about China's foreign exchange policy.

- we are able to estimate the equation for China only as opposed to panel-data sample estimations. Often, using panel-data samples researchers have been able to estimate reduced form equations for countries with few specific data. However using panel equations relies on the strong assumption of panel homogeneity and may considerably erase country specific information as mentioned by the CGER (IMF 2006) and Dunaway, Leight & Li (2009).

3 Estimation results

The series we have in hand show obvious non-stationarity in the mean (see appendix I). Empirically, these variables appear to be $I(1)$, which means that co-integration techniques can be used to test the equilibrium relationships between them (see appendix II and III for unit root tests). Using methodologies for time series with known structural breaks as in Perron (1989), Perron and Zhu (2005) and unknown structural breaks as in Zivot and Andrews (1992) we find two dates of potential structural breaks in the series of REER observations: Q1 91 and Q1 94 (see appendix III). Therefore in the remainder of this paper, we present results for the whole sample using two specifications of the deterministic component: break in Q1 91 and break in Q1 94. Also, we produce statistics for the sample starting in Q2 94, one quarter after the devaluation in Q1 94. Our first work has been to test for the cointegrating rank of the system, using Johansen's approach. As reported in table 7 in appendix IV, with a lag length of 3 for the underlying VAR, both trace and eigenvalue tests indicate one cointegrating vector at the 5% level[6]. Also, using Johansen's system approach, we check for weak exogeneity in the cointegrated system[7]. We find the joint restriction based chi-squared test marginally above the 5% critical level of rejection in the whole sample when the break is set to occur in 1991 (6.3%). The test do not reject with a trend breaking in 1994 (0.2%). This result adds a constraint in the estimation of the model, especially in the whole sample with break in 1994, because not controlling for this rejection might yield inconsistent estimators of the parameters of the equation. The rejection of weak exogeneity might be particularly binding for forecasting as it might weaken the usefulness of estimated models for economic policy. If we remove the effect of the the devaluation of Q1 1994 by testing in two subsamples, one covering Q1 1980 to Q2 1993, the second covering Q3 1994 to Q4 2008, we are able to reject the absence of weak exogeneity at 9% in the sample before the devaluation and at 27% in the sample after the devaluation. This shows that all our work, even though statistically acceptable, might be sensitive to the treatment of the one time break of 1994. Therefore, in this paper, greatest attention is paid to data manipulation and model implications.

As we are only interested by the relationship in which REER is the endogenous variable, we have reduced the system equation into a single equation as in (7) and rely on the two-step Engle-Granger (1987) procedure to estimate it. We do so, because this procedure is much more robust than the system approach to incorrect modeling of the deterministic components which is central in this paper, Hargreaves (1994)[8]. The first step of the procedure consists in estimating the long-run relationship described in equation (7), which we carry out using various corrective methodologies accounting for the proven endogeneity of the variables entering the regression. The second step consists in estimating an error correction model (ECM). We obtain a parsimonious, empirically constant, data-coherent, encompassing ECM for the Chinese REER by using a recent multi-path model selection algorithm as described in Hendry and Krolzig (2001). As inference in finite samples is commonly of concern, especially when dealing with tails of statistics distribution, we have estimated distributional properties of error correction tests by Monte Carlo simulations as in Ericsson and MacKinnon (2002). Also, this procedure enables us to take account of the common factor restriction problem in the Engle-Granger procedure as mentioned in Kremers, Ericsson, and Dolado (1992). This common factor restriction might result in a loss of power of the usual test of cointegration. We present the results of these two steps in the following sections.

3.1 Long-run relationship

We have estimated the long-run relationship described in equation (7) using OLS techniques. As we reject weak exogeneity for one specification of the deterministic component, we use the parametric correction for OLS of Stock and Watson (1993) called Dynamic OLS. This correction includes lagging and leading differences of the regressors in the equation. This may remove the effects of the simultaneity bias and autocorrelation in the residuals enabling one to obtain efficient estimators of the cointegrating vector. Standard errors are computed using the Newey and West (1987) covariance matrix with the truncation of four lags. We work in a sample of 116 quarterly observations for China only covering the period Q1 1980 to Q4 2008.

The equation to be estimated is the following:

$$y_t = \alpha + \gamma D_t + \beta X_t + \sum_{i=-q}^{q} \psi_i \Delta X_{t-i} + \mu_t \tag{8}$$

with X_t the vector of regressors at time t and D_t the deterministic component. The equation includes also lagging and leading differences of the X_t[9]. We correct the test statistics for heteroskedasticity using Newey West (1987) variance covariance estimation. The results are in table 2.

Table 2: DOLS regression statistics

Sample Q1 80 to Q4 08, trend break in Q1 91

	α	Prod	ToT	TCons	D
Estimate	2.4355	0.2601	−2.6832	2.5017	−0.0089
t value	2.0121	3.2474	−3.9664	5.3308	−9.6612
Pr(>|t|)	0.0469	0.0015	0.0001	0.0000	8.882D−16

T=111 ; R^2=0.9612 ; F(13,97)=185.271 ; DW(0)=0.6557

Sample Q1 80 to Q4 08, trend break in Q1 94

	α	Prod	ToT	TCons	D
Estimate	–	0.5190	−2.0770	2.8653	−0.0084
t value	–	8.1448	−3.3850	4.5717	−10.3957
Pr(>|t|)	–	1.244D-12	0.0010	0.0000	0.0000

T=111 ; R^2= – ; DW(0) =0.47412

Sample Q2 94 to Q4 08

	α	Prod	ToT	TCons	D
Estimate	−2.0729	0.4530	0.4731	1.2467	–
t value	−2.5705	8.0649	1.1873	6.6272	–
Pr(>|t|)	0.0134	2.373D-10	0.2411	3.307D-08	–

T=59 ; R^2= 0.8829 ; F(12,46) = 28.9182 ; DW(0) =0.681210

where α is the intercept, Prod stands for Productivity, ToT for Terms of Trade and TCons for Total Consumption as defined in the preceeding section. D is the trend.

Coefficients are all significant and with the expected right signs at the 1% level with the exception of OPEN in all samples and ToT which is non significant in the shorter sample[10]. Recursive estimations of the parameters enable us to check for their stability along the sample (see plots in appendix V). They show better stability when the structural break is set to occur in Q1 91. Also, CUSUM and CUSUM square tests do not reject stability. Visually, the coefficients seem to be invariant, at least, to smooth changes in policies and might only be sensitive to significant shocks like the devaluation in 1994. These results enable us to be fairly confident about the meaningfulness of the estimated parameters especially when the estimation is based on the Q1 91 structural break. KPSS tests as proposed by Shin (1994) tend to validate cointegration between the series[11]. However, as test statistics are commonly of concern in finite samples, we rely on the Ericsson and MacKinnon (2002) methodology to definitely test on cointegration. The final results are in the following section.

3.2 Error Correction Model

ECM estimation is the second step of the Engle-Granger procedure. Many alternative single-equation models might be obtained depending upon the model search path followed. Also, the cost of the search can be very high as the number of potential final models increases with the number of variables in the initial set, n variables implying 2^n potential models. In this paper, ECM path dependence in model selection is addressed by using the General-To-Specific approach (GeTS) developed by Hendry and Krolzig (2001). This is a multi-path model selection algorithm that improves upon Hoover and Perez's (1999) automated model-selection methodology. Ericsson (2009) presents GeTS as an algorithm which utilizes one-step and multi-step simplifications along multiple paths, diagnostic tests as additional checks on the simplified models, and encompassing tests to resolve multiple terminal models. In a four-stage process, Gets estimates and evaluates the generalized unrestricted model and checks for congruence (stage 0), simplifies it employing multi-path searches, all the while ensuring that the diagnostic tests are not rejected (stage 1), tests and simplifies union models (stage 2) and finally re-estimates the final model over two subsamples and reports the results (stage 3). If a variable is statistically significant in the full sample and in both subsamples, then the inclusion of that variable in the final model is regarded as 100% reliable.

The initial ECM to be estimated in stage 0 has the form:

$$\Delta y_t = \alpha + \sum_i \beta_i \Delta X_{t-i} + \sum_j \gamma_j \Delta Y_{t-j} + \zeta \hat{e}_{t-1} + \mu_t \tag{9}$$

with $\hat{e}_{t-1}$ the estimation of the error term of the long-term equation (8). The simplification of (9) through the four stages of the general-to-specific methodology yields the final parsimonious, empirically constant, encompassing ECM.

We successively apply this methodology to the three samples (whole sample with two different dates of potential structural break, Q1 91 and Q1 94 and the sample covering Q2 94 to Q4 08) strictly imposing the cointegrating vectors estimated by DOLS for each path tested by the algorithm. We choose to work with the liberal version of GeTS[12]. We present the results of stage (0) initial regression in table 3.

Table 3: Gets initial regression statistics

Sample Q1 80 to Q4 08, trend break in Q1 91

	α	$\hat{e}_{t-1}$	ΔY_{-1}	ΔY_{-2}	ΔY_{-3}	ΔP	ΔP_{-1}	ΔP_{-2}	ΔP_{-3}	ΔT	ΔT_{-1}	ΔT_{-2}	ΔT_{-3}	ΔC	ΔC_{-1}	ΔC_{-2}	ΔC_{-3}
Estimate	−0.00	−0.4	0.2	0.09	0.1	−0.3	0.02	0.3	−0.5	−0.7	0.8	−0.1	1.3	2.9	−1.0	−1.1	−1.6
t value	−1.3	−4.3	2.4	0.9	1.9	−0.5	0.04	0.5	−0.7	−0.8	0.8	−0.1	1.3	2.1	−0.6	−0.7	−1.0
Pr(>|t|)	0.1	0.00	0.01	0.34	0.05	0.61	0.96	0.55	0.44	0.40	0.37	0.84	0.18	0.03	0.53	0.47	0.28

T=111 ; $R^2 = 0.3447$; F(19,91)=2.519 ; DW(0)=2.0132
D-H=39.56 (0.0000) ; AR(4)=0.9547 (0.4366) ; ARCH(4)=0.2627 (0.9012)
Chow 50% = 0.2610(0.999) ; Chow 90% = 0.1680(0.998)

Sample Q1 80 to Q4 08, trend break in Q1 94

	α	$\hat{e}_{t-1}$	ΔY_{-1}	ΔY_{-2}	ΔY_{-3}	ΔP	ΔP_{-1}	ΔP_{-2}	ΔP_{-3}	ΔT	ΔT_{-1}	ΔT_{-2}	ΔT_{-3}	ΔC	ΔC_{-1}	ΔC_{-2}	ΔC_{-3}
Estimate	−0.0	−0.2	0.1	0.0	0.1	0.1	0.1	0.5	−0.3	−0.8	0.5	−0.4	1.0	3.1	−0.7	−1.5	−1.2
t value	−1.8	−3.9	1.8	0.5	1.5	0.2	0.2	0.8	−0.5	−0.8	0.5	−0.4	1.0	2.2	−0.4	−0.9	−0.7
Pr(>|t|)	0.06	0.00	0.06	0.58	0.13	0.82	0.83	0.39	0.58	0.38	0.59	0.62	0.29	0.02	0.64	0.35	0.44

T=111 ; $R^2 = 0.3290$; F(20,90)=2.206 ; DW(0)=1.9997
D-H=45.39 (0.0000) ; AR(4)=1.372 (0.2501) ; ARCH(4)=0.2278 (0.9222)
Chow 50% = 0.189(0.999) ; Chow 90% = 0.0978(0.998)

Sample Q2 94 to Q4 08

	α	$\hat{e}_{t-1}$	ΔY_{-1}	ΔY_{-2}	ΔY_{-3}	ΔP	ΔP_{-1}	ΔP_{-2}	ΔP_{-3}	ΔT	ΔT_{-1}	ΔT_{-2}	ΔT_{-3}	ΔC	ΔC_{-1}	ΔC_{-2}	ΔC_{-3}
Estimate	−0.0	−0.3	0.0	0.0	−0.0	−0.0	−0.1	−0.1	−1.0	1.2	0.2	0.2	−0.4	1.2	−0.0	−0.4	−0.8
t value	−1.4	−2.2	1.2	0.1	−0.3	−0.0	−0.1	−0.0	−0.8	2.2	0.4	0.5	−0.9	1.3	−0.0	−0.3	−0.7
Pr(>|t|)	0.1	0.0	0.2	0.9	0.7	0.9	0.8	0.9	0.3	0.0	0.6	0.6	0.3	0.1	0.9	0.6	0.4

T=59 ; $R^2 = 0.4636$; F(19,39)=1.7746 ; DW(0)=1.6926
D-H=2.0229 (0.3636) ; AR(4)=0.5945 (0.6689) ; ARCH(4)=0.220 (0.9260)
Chow 50% = 0.777(0.716) ; Chow 90% = 0.700(0.650)

We successively apply the GeTS methodology to the three samples (whole sample with two different dates of potential structural break, Q1 91 and Q1 94 and the sample covering Q2 94 to Q4 08) using the cointegrating vectors estimated by DOLS in the preceeding section and the liberal version of GeTS. For presentation purposes, we do not report coefficients for lags superior to three. As checking for congruence in the four stages of the methodology is key to generate valid results, we report a battery of diagnostic test statistics as recommended by Hendry (1995). For each sample, we report the Doornik and Hansen normality test D-H, the LM autocorrelation test with four lags AR(4), the ARCH heteroskedasticity test, the Chow predictive failure tests with break at 50% and 90% of the sample. H_0 is normality for D-H, no autocorrelation at order four for AR(4), Homoskedasticity for ARCH(4), constancy over X% of the sample for Chow.

Checking for congruence in the four stages of the methodology is key to generate valid results. We report a battery of diagnostic test statistics as recommended by Hendry (1995). Most of these statistics do not reject standard levels. Residuals appear white noise (LM and ARCH tests). They confirm that transforming yearly data to quarterly data has not generated artificial autocorrelation or heteroskedasticity. Doornik and Hansen normality tests do not reject in the short sample but do in the longer ones. That is probably because the break in the long-term series produces tail residuals that do not conform with the normal distribution. We also report the empirical constancy of the model in shorter samples (90% and 50% of the initial sample). None of the breakpoint chow statistics is significant at the 5% level. That is, no split of the sample obtains a rejection of constancy.

Then running stage 1, 2 and 3 of GeTS yields the following results for the final model in the three samples:

Table 4: Gets final regression statistics

Sample Q1 80 to Q4 08, trend break in Q1 91						
	α	$\hat{e}_{t-1}$	ΔY_{-1}			ΔC_{-4}
Estimate	−0.0093390	−0.3111052	0.2053436			1.637992
t value	−3.7631319	−5.4446054	2.41708			2.2450781
Pr(>|t|)	0.0002743	0.0000003	0.0173368			0.0268196
T=111 ; R^2 = 0.2508; F(3,107) = 11.9405						
DW(0)=1.9881 ; D-H=42.90 (0.0000) ; AR(4)=0.6973 (0.5954) ; ARCH(4)=0.2979 (0.8786)						
Chow 50% = 0.2314(0.999) ; Chow 90% = 0.3115(0.981)						
Sample Q1 80 to Q4 08, trend break in Q1 94						
	α	$\hat{e}_{t-1}$	ΔY_{-1}	ΔY_{-3}		ΔC
Estimate	−0.0065019	−0.2909796	0.2145715	0.1806074		1.5902871
t value	−2.7658433	−5.1754607	2.4897745	2.0832252		2.0555313
Pr(>|t|)	0.0066992	0.0000011	0.0143361	0.0396362		0.0422855
T=111 ; R^2 = 0.2377; F(4,106)=8.264						
DW(0)=2.0654 ; D-H=58.74 (0.0000) ; AR(4)=0.4376 (0.7811) ; ARCH(4)=0.5995 (0.6637)						
Chow 50% = 0.1331(0.999) ; Chow 90% = 0.1099(0.998)						
Sample Q2 94 to Q4 08						
	α	$\hat{e}_{t-1}$	ΔY_{-1}	ΔT	ΔC	ΔC_{-4}
Estimate	−0.0107686	−0.3000154	0.0862144	1.2429107	0.7996322	0.9951544
t value	−3.2397655	−5.0142806	1.6707362	3.5452411	1.6433564	1.9838469
Pr(>|t|)	0.0020684	0.0000063	0.1006695	0.0008292	0.1062294	0.0524629
T=59 ; R^2 = 0.4197; F(5,53)=7.669						
DW(0)=1.7158 ; D-H=0.667 (0.716) ; AR(4)=0.7457 (0.5655) ; ARCH(4)=0.9628 (0.4359)						
Chow 50% = 0.9780(0.5273) ; Chow 90% = 0.6746(0.6705)						

Again, we report the battery of diagnostic test statistics as recommended by Hendry (1995). None of these statistics reject standard levels except normality of the residual in longer samples. Residuals appear white noise (LM and ARCH). Transforming yearly data to quarterly data has not generated artificial autocorrelation or heteroskedasticity. Empirical constancy of the model in shorter samples (90% and 50%) appears acceptable. None of the breakpoint chow statistics is significant at the 5% level. Results show that the final model is parsimonious, empirically constant, data-coherent, encompassing and congruent.

In the results above, the error correction terms appear highly significant pointing to the validity of the cointegration hypothesis. However, ECM statistics' distribution is commonly of concern in finite samples, particularly in the tails. Also the usual unit root test imposes a common factor restriction on the dynamics of the relationship between the variables involved[13]. If that restriction is invalid, a loss of power may well result. To overcome these limits, we run final test of cointegration using Ericsson and MacKinnon (2002) methodology for finite sample inference about cointegration. The statistics distributional properties are estimated by Monte Carlo simulations. Three design parameters are central to the statistics' distributional properties: the sample size, the total number of variables and the number of deterministic components. The dependence of the critical values on the estimation sample size can be approximated by regression, regressing the Monte Carlo estimates of the critical values on functions of the sample size. Such regressions are response surfaces and critical values as well as p-values can be calculated from these response surfaces. Critical values are respectively -5.64, -5.27, -5.13. These compare with the t-statistics of the error terms in table 4, respectively -5.44, -5.17, -5.01. P-values are inferior to 1% therefore the null hypothesis of no-cointegration is rejected at a 1% level for the three samples.

3.3 Economic interpretation

All sample estimations point to the same conclusion: the real effective exchange rate of the Chinese Renminbi and its fundamentals have followed comparable paths during the period (see appendix V). The devaluation of 1994 might have dislocated the relationship for a while but cointegration seems to be fairly acceptable in all three samples. However more stable, hence meaningful, parameters are obtained when the structural break is set to occur in Q1 1991. As of Q4 2008, the effective exchange rate of the currency does not exhibit a significant deviation from its fair value. This result is largely in line with comparable works as Wang (2004), Funke and Rahn (2005) and Stolper and Fuentes (2007)[14].

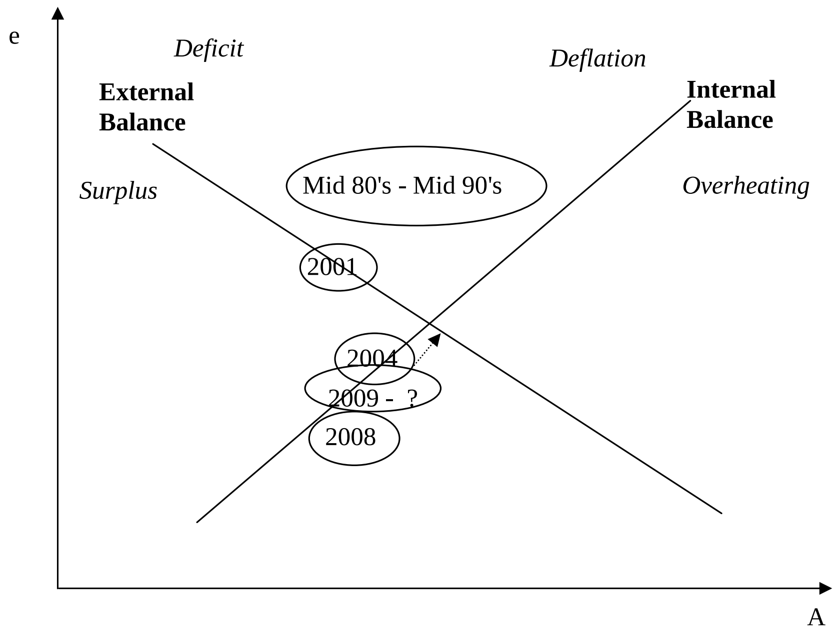
Does this conclusion close the debate about the revaluation of the currency? In this paper, we check for the time consistency of the determination of the Chinese real effective exchange rate. We do not calculate the correct value and the subsequent misalignment of the currency. By construction, the estimation of a BEER for a single currency produces small deviations from the equilibrium. Especially in the case of the RMB, which exhibits low volatility, the deviation rarely exceeds 5%[15]. If the estimation of the BEER does not enable us to conclude on the over or under valuation of the RMB, it provides strong evidence on the consistency of its determination with developments in fundamentals. This is a central result as mentioned by Edwards (1994) and advocated in many economic models; countries that maintain their REER close to equilibrium tend to overperform in

the long run. As a consequence, a sharp and immediate revaluation of the currency would surely weaken both the cointegration relationship and the development of China. Also, our main finding of time consistency in the determination of the RMB forces us to be time coherent in our analysis. Therefore we note that any call for current undervaluation implies the recognition of past overvaluation especially during episodes of deterioration of the external accounts.
In our view, the solution to the problem of the valuation of the RMB hinges on the determination of the optimal path of growth for China:

- Export-led growth calls for maintaining the current foreign exchange policy. However, this might have some implications for China especially i) in terms of inflationary pressures due to large capital inflows and ii) in terms of relationships with main trading partners.
- Inversely rebalancing growth from exports to internal consumption might be sufficient to create safe enough circumstances to weather a significant revaluation of the currency.

Especially, in line with the literature (see table 1), we find that the REER of China depends positively on productivity (Prod) and final consumption (Tcons). Therefore, our results show that, investing the proceeds of international trade domestically, mainly through government expenditure, may justify a significant appreciation of the REER. Our estimation is compatible with the mid-term version of the natural real exchange rate (NATREX) of Stein (1994). Like BEER, NATREX is a positive methodology. It does not impose norms on fundamentals to estimate the fair value of the currency. However, it imposes a stricter framework of analysis in which technological progress, terms of trade and time preferences are key factors[16]. This might have important implications as the NATREX is defined as the intercyclical real exchange rate that ensures the balance of payments' equilibrium in the absence of cyclical factors, speculative capital movements and changes in monetary policy. Since it is an equilibrium concept, tracking the NATREX should guarantee the attainment of both internal and external equilibriums in the long run. Therefore, maintaining the current relationship between the currency and its long-term fundamentals as estimated in this paper should enable China to move safely along any chosen optimal path of growth. We have estimated our model while accepting weak exogeneity at the margin which means that great attention must be paid to the meaningfulness of the coefficients. Especially, as mentioned by Baffes, Elbadawi and O'Connell (1997), the relevant issue is whether the coefficients might be treated as a constant in the face of shifts in the marginal distribution of the fundamentals as to control for Lucas' critique of econometric policy. Impulse response analyses is often an answer to these questions, but they require careful handling especially in a rather small sample. In our sample, they exibit, even though permanent, really negligible reactions of size around 0.02%. But we prefer to rely on the observation of the behaviour of the coefficients which indicates that our model estimation seems fairly stable when facing smooth changes in fundamentals[17]. Therefore, it seems suitable to clarify what the common behaviour of the REER and its fundamentals should be to remain on the equilibrium path determined by the NATREX[18]. Using the results of the long-term relationships (see table 2), we estimate that a shock of 1% on absorption might justify an adjustment of the REER, in the same direction, by around 2.5% in the long run. This conclusion stands for the long run and is not related to the current economic conditions of China. From a graphical standpoint, such a policy would draw China closer to the equilibrium point of the Swan (1963) diagram. Final consumption might increase thanks to the development of a public social security networks that in turn might push the consumer saving rates lower. Also investment in the banking system, especially restructuring state banks' non-performing loans might push private consumption higher. Finally, final consumption might increase thanks to investment in the education system. This appears particularly important, especially for productivity, in a country where around 20 to 25 million new urban job-seekers come on the market every year.

Figure 1: Swan diagram for China



e = real exchange rate (a rise means an appreciation) and A is absorption

This diagram, Swan (1963) provides a simple framework for understanding the relationship between the REER and internal and external balance. The upward-sloping curve is for "internal balance". In the top left region the country suffers from unemployment and / or deflation. The downward-sloping curve is for "external balance". In the bottom left region, the country's balance of payment exhibits current account surplus. The circles situate the economy in the diagram for the corresponding year or period. Increasing absorption would push China further to the top right region along the internal balance line.

4 conclusion

The Chinese economy exhibits impressive external accounts: as of end 2009, the surplus of the current account represents around 10% of the GDP of the country while international reserves amount to around USD 2000bn. This is more than twice the cumulated reserves of the three largest holders behind China[19]. And one has to cumulate the holdings of the forty largest emerging countries to build reserves equivalent to those of China. Trade surpluses and foreign exchange reserves accumulation have put the RMB under verbal pressure around the world both in the academic and in the business communities. Negotiating some sort of shadow limits for the CA surplus of China seems to be on top of the international political agenda. According to many estimations, reducing the CA surplus to more acceptable levels would imply a revaluation of the currency ranging from 20% to 40%.
In this paper, we produce helpful results to think about any policy aiming at reducing current imbalances. Especially, using recent developments in cointegration methodologies, we give evidences on the time consistency of the determination of the Chinese Renminbi since the early 1980s. We show that, in line with the natural exchange rate of Stein (1994), the behaviour of the currency is intimately related to such fundamentals as productivity, absorption and the terms of trades. Using an appropriate econometric method and a proper specification, we found results that are data-coherent, empirically constant and robust to the most recent and harshest diagnostic tests. Our results show that an increase of absorption over the coming years should create safe enough domestic circumstances to weather a significant revaluation of the currency in the long run. Also, they enable us to represent the concomitant internal and external equilibrium in a REER-Absorption space as in the Swan diagram. In the framework of the Swan diagram, our results confirm that increasing absorption would push China further to the top right along the internal equilibrium line. Such a strategy would have the advantage to seek external equilibrium while protecting domestic growth which seems particularly important in a country that has to deal with large numbers of unemployed persons. Our results may have important implications in term of economic policy as they enable one to quantify precisely the shock on absorption that is needed to reach any position in the diagram. Especially, we estimate that any increase over the long term of 1% of the ratio of absorption to GDP might justify a revaluation of 2,5% of the real effective exchange rate of China.

Our work has little to say about the correct value of the currency: the target level of the REER depends on the determination of the optimal path of growth for China. However, our main finding of time consistency in the determination of the RMB forces us to be time coherent in our analysis. Therefore any call for current undervaluation implies the recognition of past overvaluation especially during episodes of deterioration of the external accounts. Edwards (1994) reports that countries that have maintained their REER close to the equilibrium have systematically over performed. Therefore, whatever the decision of the Chinese authorities, the gradual convergence to the targeted equilibrium exchange rate must be in line with developments in fundamentals. This is one of the key conditions for maintaining the economy on an optimal path.

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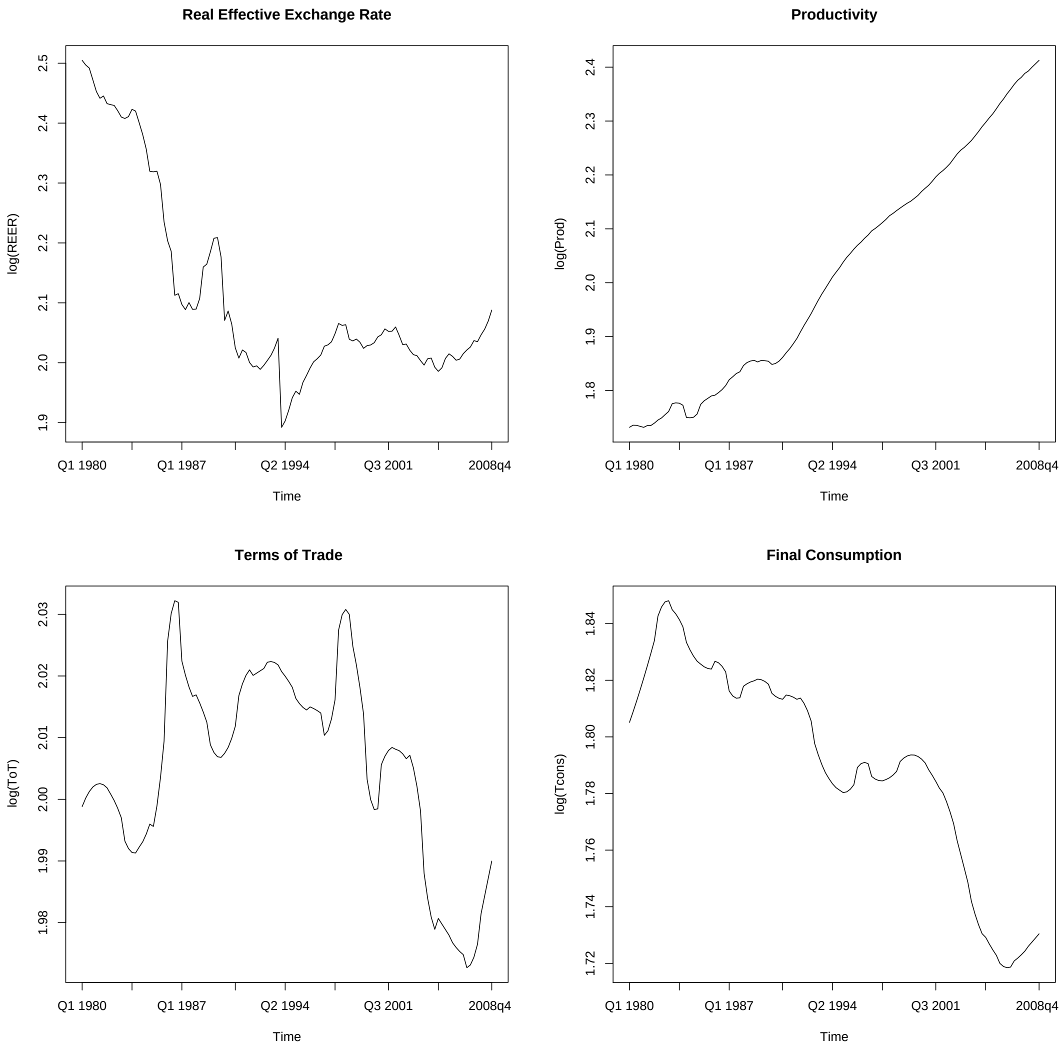
Appendix I

Times Series Plots



The series we have in hand show an obvious non-stationarity in the mean, being seemingly integrated of order 1, i.e. I(1). The series of Real Effective Exchange Rate exhibits a change in the trend that we must account for. Especially, the break in the REER series seems to correspond to the devaluation that occurred in January 1994. Given the presence of a break in the series that might affect the results of the statistical analysis, we have estimated the model in two samples: in a sample starting in Q1 80 and in a sample starting in Q2 94. In appendix II we test for unit roots in the post devaluation sample starting in Q2 1994. In appendix III, we test for unit roots in the whole sample starting in Q1 1980.

Appendix II

Pre-testing for unit roots is the first step in the cointegration modeling. Especially, determining the most appropriate form of the trend in the data is key to producing acceptable empirical results. There exists an abundant literature on unit root tests starting with the now very popular tests developed by Dickey and Fuller (1979 and 1981) to more recent and powerful tests. Due to its wide diffusion and simplicity we start with the ADF test of Dickey and Fuller and then adjust step by step to more powerful tests. Especially we perform the Phillips-Perron test (hereafter PP) to account for serial correlation and heteroskedasticity in the errors, the Schmidt-Phillips test (hereafter SP) to deal with the uncertainty about the form of the trend in the series and Elliott-Rothenberg-Stock test (hereafter ERS) to control for near unity loading in the ADF regression. Basically all these tests use the procedure of the ADF test with the null hypothesis of nonstationarity but either with filtered data (SP, ERS), adjusted statistics (PP) or slightly different specifications in the regression (SP). Finally we also perform the KPSS test that has cointegration as null hypothesis to confirm the results. In table 5, we produce unit root tests for the original variables (all in logs) in the sample Q2 94 to Q4 08.

Table 5: Unit Root test: 94 Q1 - 08 Q4

	ADF $<$ -3.58	PP $<$ -3.58	SP $<$ -25.2	ERS $<$ -2.58	KPSS $>$ 0.73
REER	-3.0	-2.6	-5.1	0.1	0.8
Prod	0.7	-0.6	-9.1	1.0	3.0
TOTW	-1.0	-1.5	-16.9	-1.2	2.5
OPEN	0.8	-1.8	-6.2	0.4	2.8
OPENM	0.6	-1.9	-7.9	-0.2	2.7
GCON	-1.2	-1.7	-19.2	-2.0	0.5
GPCON	0.3	-1.4	-8.7	-0.9	2.5
NFA	-0.2	-2.5	-17.3	1.4	3.0

H_0 is non-stationarity for the ADF, PP, SP, ERS tests. H_0 is stationarity for the KPSS test. We provide 1% critical values for each test.

The results of these tests confirm the visual impression: non-stationarity is not rejected with a risk of 1% for all tests and all variables with the only exception of the ADF test for REER which does reject non-stationarity. Also using KPSS test, we can reject stationarity for all variables. Therefore we can reasonably conclude that the series are $I(1)$.

An important practical issue for the implementation of the ADF test is the specification of the lag length p that sets the order of serial correlation and heteroskedasticity of the series. If p is too small the remaining serial correlation in the errors will bias the test. If p is too large then the power of the test will suffer. We have first set an upper boundary of p=4 and estimate the regression. Minimizing the Akaike Information Criterion (AIC) and controlling for significance of loadings we have shortened the lag horizon to one for many series. With p=1 the 1% critical value for the ADF test is -3.58. The same test procedure is applied, when appropriate, for other tests.

Appendix III

The series of Real Effective Exchange Rates (REER) shows an obvious structural break in the trend in the early 1990s. This might correspond to the devaluation of the Renminbi in January 1994. Not accounting for this break in the series would largely affect the results of the estimation of the model using co-integration techniques. Especially, the series of residuals would be heteroskedastic affecting test statistics. Following Perron (1989) we correct for this structural break by introducing the variable Tb whose value is $Tb_t = t - Tb$ if $t < Tb$ and 0 otherwise, with Tb the date of the structural break.
Two approaches coexist to determine the date of the structural break: either the break is exogenous and corresponds exactly to a certain event in history, in this sample for instance the Q1 94 devaluation, or the break is endogenous and chosen to fit a certain representation of the series. Perron (1989) develops a procedure for testing the null hypothesis that Y_t has a unit root given that an exogenous structural break occurs at time Tb. This is tested versus the alternative hypothesis that the series is stationary around a deterministic time trend with an exogenous change in the trend function at time Tb. Perron's parameterizations of the structural break are the following:

$$y_t = \mu_1 + \ y_{t-1} + (\mu_2 - \mu_1)DU_t + \epsilon_t \tag{10}$$

Where $DU_t = 1$ if $t > Tb$, 0 otherwise. The trend stationary alternative is:

$$y_t = \mu_1 + \beta_1 t + (\beta_2 - \beta_1)DT'_t + \epsilon_t \tag{11}$$

Where $DT'_t = t - Tb$ if $t < Tb$, 0 otherwise.

Then his test for a unit root involves the following augmented regression equations:

$$y_t = \mu_1 + \beta_1 t + \omega_1 DT'_t + \alpha y_{t-1} + \sum_{j=1}^{k} \psi_j \delta y_{t-j} + \epsilon_t \tag{12}$$

To formally test for the presence of a unit root, Perron considers the standard t-statistic for testing $\alpha = 1$

Performing Perron's test, with structural break in Q1 94, we cannot reject the null of unit root for REER in the sample from Q1 80 to Q4 08. Perron has estimated critical values for his test. Results for the sample Q1 80 to Q4 08 are in table 6 where we also present unit root test results for all series in the sample. The tests confirm that all series are not stationary.

Following Perron's approach, Zivot & Andrews (1992) question the exogeneity of the structural break and treat it as an endogenous occurrence. Doing so, they view the selection of the unknown break point, $T\omega$, as the outcome of an estimation procedure designed to fit y_t to a certain trend stationary representation. That is they assume that the alternative hypothesis stipulates that y_t can be represented by a trend stationary process with a one time break in the trend occurring at an unknown point in time. The ADF test equation becomes:

$$y_t = \mu + \beta_t + \lambda DT'_t(\omega) + \alpha y_{t-1} + \sum_{j=1}^{k} \psi_j \delta y_{t-j} + \epsilon_t \tag{13}$$

Where $DT'_t(\omega) = t - [T\omega]$ if $t > [T\omega]$, 0 otherwise
Basically, rolling the break point along the sample, they choose the break point date that gives

Table 6: Unit Root test: 80 Q1 - 08 Q4

	ADF < -3.58*	PP < -3.58	SP < -25.2	ERS < -2.58	KPSS > 0.73
REER	−2.7	−1.2	−5.4	−0.2	3.8
Prod	3.5	−2.6	−8.0	2.4	5.8
TOTW	−0.9	−1.7	−17.1	−2.1	1.6
OPEN	−1.8	−3.4	−32.7	1.2	5.6
OPENM	−2.0	−3.1	−37.9	0.7	5.3
GCON	−1.5	−1.9	−35.6	−2.7	0.4
GPCON	1.3	−3.1	−17.0	−0.7	5.0
NFA	−0.9	−3.0	−19.2	1.4	5.5

*Perron has estimated critical values for his test. The critical value for ADF unit root test with a known structural break for REER is -4.6 at 1%. Other series refer to usual ADF critical values, -3.58 at 1%.

the least favorable result for the null hypothesis. Using the procedure described in Appendix II we select an optimal lag of 0. We find a significant structural break point in the trend occurring in Q1 91. The test statistic is -3.11 while the 1% critical value is -4.93. Therefore, using Zivot & Andrews as Perron's methodology, does not enable one to reject the presence of a unit root in the sample Q1 80 to Q4 08.

Appendix IV

Table 7: Johansen's trace and eigenvalue test of cointegration rank

Sample Q1 80 to Q4 08, break in 91				
null	Trace stat	90%	95%	99%
r $\leqslant$ 0	92.87	79.82	84.66	94.37
r $\leqslant$ 1	47.53	53.95	58.03	66.31
null	ME stat	90%	95%	99%
l $\leqslant$ 0	45.33	37.87	41.07	47.73
l $\leqslant$ 1	29.14	30.46	33.44	39.66

Row one and two test respectively the null hypothesis of no cointegration and the null hypothesis of at most one cointegrating vector. The second column reports the trace statistic while column three, four and five report small-sample-adjusted 10%, 5% and 1% critical values for trace test as in Cheung and Lai (1993). The same applies for Max Eigenvalue test (ME stat) at line four and five. We test with 3 lags for the VAR which includes the four endogenous variables, a trend breaking in Q1 1991 and a dummy with value 0 from Q1 1980 to Q4 1993 and then 1.

Table 8: Johansen's trace and eigenvalue test of cointegration rank

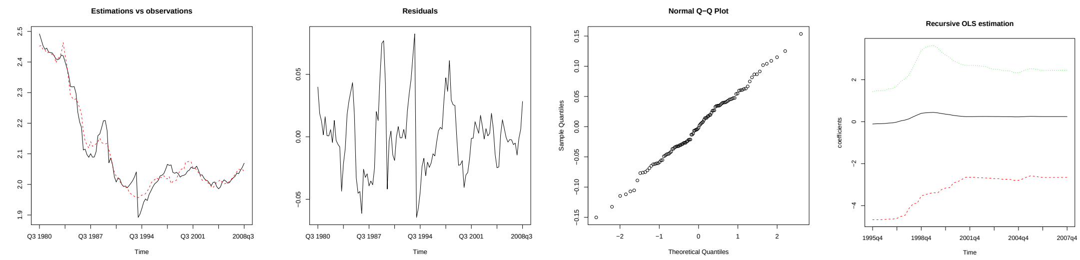
Sample Q1 80 to Q4 08, break in 94				
null	Trace stat	90%	95%	99%
r $\leqslant$ 0	84.29	79.82	84.66	94.37
r $\leqslant$ 1	48.88	53.95	58.03	66.31
null	ME stat	90%	95%	99%
l $\leqslant$ 0	35.41	37.87	41.07	47.73
l $\leqslant$ 1	26.45	30.46	33.44	39.66

Row one and two test respectively the null hypothesis of no cointegration and the null hypothesis of at most one cointegrating vector. The second column reports the trace statistic while column three, four and five report small-sample-adjusted 10%, 5% and 1% critical values for trace test as in Cheung and Lai (1993). The same applies for Max Eigenvalue test (ME stat) at line four and five. We test with 3 lags for the VAR which includes the four endogenous variables and a trend breaking in Q1 1994.

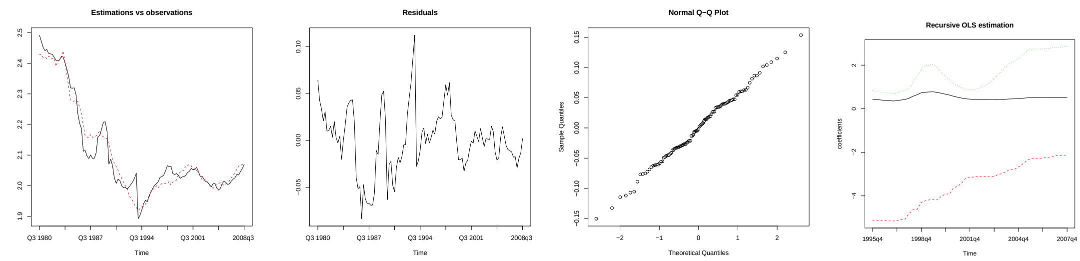
Appendix V

Estimations, Observations and Residuals

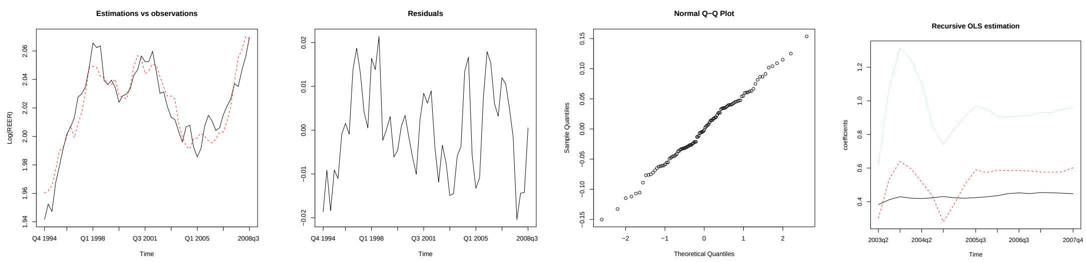
Sample Q1 80 to Q4 08, break in Q1 91



Sample Q1 80 to Q4 08, break in Q1 94



Sample Q2 94 to Q4 08



Notes

[1]This equation holds if fundamentals are stationary in the first difference (i.e. I(1)).

[2]Models implied by the macro economic balance approach like Faruqee (1995) or Stein (1994) share most of the fundamentals cited by Elbadawi (1994). Especially they all share a saving-absorption bias and a Balassa-Samuelson bias. Also terms of trade appear central in all models.

[3]See Egert (2004) and Lee, Milesi-Ferretti and Ricci (2008).

[4]Bertsekas (1976). We used a quadratic average which fits a local quadratic polynomial for each observation of the yearly frequency series then used this polynomial to fill in all observations of the quarterly series.

[5]See Ferruci(2003) and Dell Ariccia & al. (2006).

[6]both trend specifications enable us to accept, at least, one cointegrating vector at the 5% level. The test based on a trend breaking in Q1 1991 includes also a dummy variable with value 0 from Q1 1980 to Q4 1993 and then 1.

[7]Weak exogeneity holds if the cointegrating vector does not enter the marginal model for the fundamentals, see Urbain (1992). We test this hypothesis by constraining the coefficients of the error terms in the marginal models to be zero.

[8]Also the Engle-Granger method is more robust to the presence of I(2) variable, to processes with fractional unit roots and to processes with stochastic unit roots as mentioned by Gonzalo and Lee (1998). Also, it is well known that Johansen procedures deteriorate in small samples due to the number of parameters to be estimated.

[9]The specification of the lag and lead lengths is comparable to the specification procedure used for unit root tests. This procedure is described in Appendix II.

[10]the intercept might be adjusted for the Penn effect. However, this effect has no consequence on the other parameters of the equation, hence on the conclusions of this work.

[11]The KPSS statistics are respectively 0.1462, 0.1238 and 0.1235 while the 10% critical value is 0.347.

[12]The general-to-specific method has been programmed in the econometric toolbox GROCER on Scilab under the name "automatic" available at http://dubois.ensae.net. The general-to-specific approach as the function "'automatic"' offers two options called "'liberal"' and "'conservative"'. As mentioned by Ericsson (2009), the liberal strategy errs on the side of keeping some variables even though they may not actually matter. Inversely, the conservative strategy keeps only variables that are clearly significant statistically, erring in the direction of excluding some variables, even though those variables may matter.

[13]See Ericsson and MacKinnon (2002) for a survey.

[14]It is different to Benassy Quéré and al. (2006) who find an undervaluation of around 40%. However this study estimates the REER for China on a large panel, including developed countries like the US and the UK and does not explore the question of break in the deterministic component.

[15]This echoes Cheung, Chinn and Fujii (2007) who mention that due to sampling uncertainty, in many empirical works, the deviation of the RMB, even though sometimes large, is not statistically significant

[16]It is usual to use productivity and absorption as respective proxies for technological progress and time preferences.

[17]As often in econometrics, extending this conclusion to large shifts might be unwise. Therefore, we do not use our estimation to simulate the behaviour of the cointegration relationship facing a significant shift in policy like for instance the introduction of floating exchange rates

[18]The rejection of weak exogeneity would imply the existence of feedback effects in the system which might result in misleading simulations. However, the feedback effect on Absorption is really small as a deviation of the REER of 1% creates only a negligible but positive effect of 0.02% on A, over the long term. Finally any work using a VECM in a small sample must be taken with great care given the high number of parameters to be estimated

[19]Namely Russia, USD 411bn, India USD 240bn and Brazil USD 192bn. The aggregation of these three countries yields a virtual economic region that would have 110% of the population of China, 90% of its nominal GDP in dollar terms but less than 50% of its foreign exchange reserves.