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The Equilibrium Real Exchange Rate of China: A Productivity Approach

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Abstract

A large body of theoretical and empirical works asserts that exchange rates depend upon a country's productivity growth, and this effect is dubbed the Balassa-Samuelson effect. This paper examines the evidence for a Balassa-Samuelson based explanation for the real exchange rate movements of China vis-a-vis the U.S. dollar. Using disaggregated industry level data, we construct sectoral total factor productivities (TFPs) for the tradable and nontradable sectors from 1980-2003. Our main findings are: (a) the sectoral TFP differential is cointegrated with the relative price of nontradables with the unit cointegration vector; and (b) the real exchange rate is cointegrated with home and foreign sectoral TFP differentials. This productivity based real exchange rate model is then used to estimate the equilibrium exchange rates of renminbi (RMB). A comparison of the equilibrium exchange rate predicted by the productivity-based model and the actual rate indicates that the Renminbi is somewhat undervalued against the US dollar, though the undervaluation is not statistically significant. Our finding is broadly consistent with the conclusion of Chang and Shao (2004), and it also suggests that the extent of Renminbi undervaluation may have been exaggerated in the recent debate. Our conclusions continue to hold even after we have controlled for the movements of net foreign assets.

Keywords: Nontraded Goods, Balassa-Samuelson Model, Cointegration

JEL Classification System: F31, F41

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1. Introduction

Of the several competing explanations for the persistent deviations of nominal exchange rates from their Purchasing Power Parities (PPPs), perhaps the earliest and most fundamental is the productivity differential hypothesis proposed by Balassa (1964) and Samuelson (1964). The Balassa-Samuelson hypothesis asserts that different trends in tradable and nontradable sectors' productivity cause systematic departures of exchange rates from PPPs by changing the relative price of nontradable (to tradable) goods. This paper examines whether the Balassa-Samuelson model can explain the real exchange rate[1] movements for China vis-a-vis the U.S. dollar.[2]

Given that the Balassa-Samuelson (henceforth B-S) model is the pre-eminent model for explaining long run real exchange rate movements, we believe it is important to examine to what extent it is successful in explaining China's real exchange rate. Since the B-S model relies on differential productivity growth rates, we would a-priori expect it to be especially relevant for determining the real exchange rates of the relatively fast growing economies like China. However, investigations of the B-S effect for China have been hampered by lack of availability of the relevant data, and despite years of rapid industrial growth, clear evidence of a B-S effect has been scant. Also, one of the attractive properties of the B-S model from a theoretical perspective is that the link between productivity differentials and the real exchange rate is robust to the underlying assumptions that are often made for expositional convenience. Obstfeld and Rogoff (1996) emphasize this robustness of the model and show that several assumptions of the model can be relaxed without affecting the key predictions of the model. Hence the B-S model likely has greater applicability to a developing economy like China than has perhaps been recognized in the extant literature.

Ito, Isard and Symansky (1996) document a positive correlation between growth rates (relative to the U.S.) and real exchange rate appreciation for a group of East Asian economies including China, but find that changes in the relative price of nontradables do not uniformly play an important role in explaining the evolution of the real exchange rate. However, a country's growth rate relative to the U.S. is at best a crude proxy for the productivity differential between the tradable and nontradable sectors. Another proxy commonly used in the literature is the ratio of consumer price index (CPI) to the producer price index (PPI) to measure the relative price of price of nontradable to tradable

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[1]The real exchange rate is defined as the ratio of the domestic price level to the foreign price level adjusted by the nominal exchange rate. With this definition, deviations of nominal exchange rates from PPP are synonymous with changes in the real exchange rate.

[2]It would also be of interest to examine the determinants of the multilateral real exchange rate of China as in Hua (2007). However, availability of industry level data required to construct the total factor productivity and ensuring comparability of our results to the existing literature constrain us to focus on the bilateral real exchange rate against the U.S.

goods. This is justified by the argument that the CPI has a relatively greater share of nontradables whereas the PPI has a higher share of tradables. Using a slightly different interpretation, Coudert and Couharde (2007) examine the evidence for the Balassa-Samuelson effect in China and other developing economies using the ratio of CPI to PPI as the proxy for the overall price level relative to the price of tradables. They do find evidence of a Balassa-Samuelson effect for most countries, but not for China.[3] However, as pointed out recently by Chinn (2005) and Égert, Lommatzsch and Lahrèche-Révil (2006), a key shortcoming of using relative prices is that they could also be influenced by demand-side factors and hence may plausibly capture effects other than the B-S effect.

Chinn (2000) constructs estimates of labor productivity in the tradable and nontradable sectors for China and other Asian countries, where the tradable sector is proxied by the manufacturing, mining and agriculture sectors and the nontradables sector comprises the rest of GDP. He finds evidence for a significant link between real exchange rates and labor productivity differentials for Japan, Malaysia and the Philippines, but not for other countries including China. More recently, however, Cheung, Chinn and Fujii (2007), extend Chinn's (2000) analysis of China's real exchange rate vis-a-vis the U.S. dollar. They uncover more favorable evidence for the Balassa-Samuelson hypothesis for China in that the Chinese and U.S. labor productivity differentials are significant determinants of Yuan-Dollar real exchange rate.[4] The importance of real shock as a source of real exchange rate fluctuations in China has been forcefully emphasized by Wang (2005). Using structural vector autoregression model, Wang finds that supply shocks account for a substantial part of the long-term variations in the real exchange rate changes of China during 1980-2003. Compared with studies on industrial countries with flexible exchange rate systems by Eichenbaum and Evans (1995) and Clarida and Gal (1994), Wang finds that supply shocks play a more important role in China due to supply-side changes from structural reforms and productivity shocks.

As the differential evolution of productivity in the tradable and nontradable sectors lies at the heart of the Balassa-Samuelson theorem, the importance of using a carefully measured productivity variable in the empirical analysis cannot be overemphasized. The main contribution of this paper is examining the evidence for the B-S effect in China by constructing measures of sectoral total factor productivity. Total factor productivity (TFP) is more consistent with the theory underlying the B-S model as it can help isolate the impact of supply-side effects. The relative price of nontradables and labor productivity variables, by contrast, may also be influenced by demand-side effects and hence

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[3] See also Funke and Rahn (2005)

[4] Cheung, Chinn and Fujii (2007) also provide a good overview of the recent literature on the Chinese real exchange rate against the U.S. dollar with a special emphasis on the Renminbi misalignment debate.

cannot accurately reveal the existence and magnitude of the B-S effect. However, measurement of TFP requires sectoral data on capital stocks, which are generally unavailable for China even at the aggregate level. We use the aggregate series for China's capital stock constructed recently by Holz (2006) and also construct our own estimate of the aggregate capital stock using investment data as a sensitivity check. The aggregate capital stock is then allocated to the tradable and nontradable sectors in proportion to the share of capital income in that sector. TFPs for each sector are computed as Solow residuals from sectoral Cobb-Douglas production functions.

Our results confirm the existence of a strong B-S effect in China. We find that the relative price of nontradables in China is driven by the TFP differential between the tradable and nontradables sectors. We also find that different trends in sectoral TFPs in the U.S. and China can successfully account for the long run changes in the China-U.S. bilateral real exchange rate. In addition to sectoral TFP differentials, we also examine the role of net foreign assets in the determination of the China-U.S. real exchange rate. Consistent with prior empirical work (Lane and Milesi-Ferreti (2004); Égert, Lommatzsch and Lahrèche-Révil (2006)), we find that an increase in net foreign assets leads to an appreciation of the Chinese real exchange rate in the long run.

With China playing an increasingly important role in global financial markets, the exchange rate policy of China has gained growing attention around the globe. One particularly important issue, which has generated intense debate among academics, policy analysts and trade officials, is whether China's currency is undervalued relative to some theoretically determined equilibrium value. Since there are several different approaches to modeling equilibrium real exchange rates, the results of any one approach must necessarily be viewed with caution (Égert and Halpern (2006)). Using the B-S model as a benchmark, we find that although there have been periods of statistically significant misalignment in our sample, the China-U.S. real exchange rate is close to its equilibrium value as at the end of 2003. Other studies on China that reach a similar conclusion include Chou and Shih (1998), which uses a purchasing power parity (PPP) approach, and Zhang (2001), which uses a behavioral equilibrium exchange rate (BEER) approach. Chang and Shao (2004), using cross-country analysis on 160 countries, similarly conclude that the real exchange rate of the Chinese currency RMB was not severely undervalued from a statistical point of view over our sample period.

The rest of the paper is organized as follows. Section 2 reviews the B-S model and derives the key relationships between the relative price of nontradables, sectoral TFP differentials and real exchange rates in a cointegration framework. Section 3 explains the econometric framework used in our empirical analysis. Section 4 presents the empirical results. Section 5 examines the robustness

of the productivity-based cointegration model by augmenting it with net foreign assets. Section 6 concludes. Details of data sources and construction are provided in appendices 1 and 2.

2. The Model

2.1. The Relative Price of Nontradables

The model used is a variant of Rogoff (1996). Each country is divided into tradable and nontradable goods sectors: good T is tradable and good N is nontradable. The production side of the economy is summarized by the following Cobb-Douglas production functions:

$$Y_T = A_T(L_T)^{\alpha_T}(K_T)^{(1-\alpha_T)}, \tag{1}$$

$$Y_N = A_N(L_N)^{\alpha_N}(K_N)^{(1-\alpha_N)}. \tag{2}$$

Here Y denotes output; L and K denote labor and capital, respectively; A denotes TFP and α denotes the share of labor in production. Country superscripts are suppressed unless it is necessary to distinguish between home and foreign countries.

Standard assumptions of the Balassa-Samuelson model yield the following set of first-order conditions:

$$A_T(1-\alpha_T)(k_T)^{(-\alpha_T)} = r = QA_N(1-\alpha_N)(k_N)^{(-\alpha_N)}, \tag{3}$$

$$A_T\alpha_T(k_T)^{(1-\alpha_T)} = w = QA_N\alpha_N(k_N)^{(1-\alpha_N)}. \tag{4}$$

Here r denotes the world real interest rate, which is determined in the world capital market; w denotes the real wage rate; k_T and k_N denote capital-labor ratios in the tradable and nontradable goods sectors, respectively; and Q denotes the relative price of the nontradable good in terms of the tradable good. The tradable good is chosen to be the numeraire good, so that the real wage rate and the real interest rate are both measured in terms of tradables.

Relation (3) equates the marginal product of capital in each sector to the world real interest rate in terms of tradables, whereas Relation (4) equates the marginal product of labor in each sector to the real wage rate in terms of tradables. Since each competitive firm takes as given the world real interest rate r, the left-hand-side equation of (3) determines the capital-labor ratio in the tradable goods sector (k_T). Given k_T, the left-hand-side equation of (4) determines the real wage rate. Given the interest rate and the wage rate, the right-hand-side equations in (3) and (4) jointly determine the relative price of nontraded-goods (Q) and the capital-labor ratio in the nontradable goods sector (k_N).

Solving for the relative price of nontradables in terms of the sectoral TFPs and the world real interest rate and taking logs yields[5]:

$$\ln(Q) = \lambda + \frac{\alpha_N}{\alpha_T} \ln(A_T) - \ln(A_N) + \frac{(\alpha_T - \alpha_N)}{\alpha_T} \ln(r). \tag{5}$$

Here λ is a constant that depends on labor shares. Equation (5) embodies the first key prediction of the Balassa-Samuelson model, which is that the relative price of nontradables within each country depends on the labor-share adjusted sectoral TFP differential and the exogenous world real interest rate.[6]

In order to derive testable implications of Equation (5) in a cointegration framework, we assume that:

Assumption 1: The relative price of nontradables is a difference stationary or I(1) process, possibly with non-zero drift.

Assumption 2: The labor-share adjusted sectoral TFP differential is a difference stationary or I(1) process, possibly with non-zero drift.

Assumption 3: $\ln(r)$ is a stationary or I(0) process.

The first assumption states the necessary condition for the relative price of nontradables to explain the stochastic trend observed in real exchange rates. If the relative prices of nontradables is stationary, the Balassa-Samuelson cannot rationalize the observed unit root behavior of real exchange rates. The second assumption states the necessary condition for the labor-share adjusted sectoral TFP differential to explain for the stochastic trend in the relative price of nontradables. It requires that technological progress is not transmitted from one sector to the other in such a way as to render the labor-share adjusted TFP differential stationary. If $\ln(A_T)$ and $\ln(A_N)$ are cointegrated with the normalized cointegrating vector $(1, -\alpha_T/\alpha_N)'$, then their linear combination would be stationary. In this special case, the stochastic trends in sectoral TFPs annihilate each other, and the Balassa-Samuelson model cannot rationalize the stochastic trend in the relative price of nontraded-goods. Assumption 2 rules out this possibility. Section 4 presents evidence supporting these two assumptions.

The third assumption is maintained for two reasons. First, most economic models predict the real interest rate to be stationary. Second, it is difficult to construct an accurate proxy for the (long term, ex ante) world real interest rate in terms of tradables. Most studies of the Balassa-Samuelson model make the stronger assumption that the world real interest rate is constant. Given that empirical work

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[5] See Obstfeld and Rogoff (1996), Chapter 4, or Kakkar (2002).

[6] Obstfeld and Rogoff (1996) show that this key prediction of the B-S model holds even without the assumptions of perfect capital mobility or two factors of production.

generally does not find any significant impact of the real interest rate on investment, the measurement error introduced in estimating the (long term, ex ante) world real interest rate will likely outweigh the benefit of extra information.

Taken together, these three assumptions imply that ln(Q) should be unit root nonstationary and cointegrated with the labor-share-adjusted sectoral TFP differential $d = (\alpha_N/\alpha_T)\ln(A_T)-\ln(A_N)$ with the normalized cointegrating vector $(1, -1)'$. The following cointegrating regression is estimated for both China and the U.S. to test whether this implication of the model is supported empirically:

$$\ln(Q) = \lambda + \delta d + \varphi. \tag{6}$$

Here φ is a zero-mean stationary random variable that captures any short run deviation of the relative price of nontradables from its long run equilibrium value. We turn next to the relationship between the relative price of nontradables, sectoral productivity differentials and the real exchange rate.

2.2. The Real Exchange Rate

Consider a world economy with two countries – country H is the home country and country F is the foreign country. We assume that the price level of each country P^i (i = H, F) can be approximated by a geometric average of the prices of nontradable and tradable goods up to a stationary measurement error:

$$P^i = c^i(P_N^i)^{\beta^i}(P_T^i)^{1-\beta^i}. \tag{7}$$

Here β^i is a constant which measures the share of nontradables in the overall price level of country i and c^i is a stationary measurement error that reflects factors which cause the general price level to deviate from the geometric average of the price of nontradable and tradable goods. Let E denote the nominal exchange rate between the home and foreign countries – E units of the home country's currency buy 1 unit of the foreign currency. The real exchange rate between the home and foreign countries, E^r, is the ratio of the home price level to the foreign price level adjusted by the nominal exchange rate:

$$E^r = \frac{P^H}{E \cdot P^F}. \tag{8}$$

It is assumed that the law of one price holds for tradable goods in the long run.

<u>Assumption 4</u>: The real exchange rate for tradable goods, $(P_T^H/E \cdot P_T^F)$,is a stationary or I(0) process.

Mathematically, we can write this assumption as

$$\ln\left(P_T^H\right) = \ln(E) + \ln\left(P_T^F\right) + u, \tag{9}$$

where u is a stationary random variable. The stationarity of u ensures that deviations from PPP for tradable goods are short lived. Equations (7)-(9) imply that

$$\ln\left(E^r\right) = \theta + \beta^H \ln\left(Q^H\right) - \beta^F \ln\left(Q^F\right) + \epsilon, \tag{10}$$

where θ is a constant and ϵ is a zero-mean stationary random variable. Equation (9) shows that the real exchange rate depends on the relative price of nontradables in the home and foreign countries. To highlight the connection between real exchange rates and TFP differentials, we combine equations (6) and (10) to get

$$\ln\left(E^r\right) = \mu + \beta^H d^H - \beta^F d^F + \eta, \tag{11}$$

where μ is a constant and η a zero-mean stationary random variable. Equation (11) is the crux of the Balassa-Samuelson model as it implies that the real exchange rate is determined by the home and foreign sectoral TFP differentials in the long run. An increase in the home sectoral TFP differential, which means faster TFP growth in the tradable sector relative to the nontradable sector, is associated with a higher relative price of nontradables via equation (6) and an appreciating real exchange rate via equation (11).

Equations (6) and (11) are the key testable predictions of the Balassa-Samuelson model and form the basis of the empirical work. Taking China to be the home country and the U.S. as the foreign country, we employ the canonical cointegrating regression (CCR) procedure to estimate these two equations. Since there is a possible structural break in the real exchange rate when China reformed its exchange rate system by unifying its dual exchange rates at the end of 1993[7], we apply Gregory-Hansen's (1996) test to detect the presence of a structural break in the intercept and/or slope coefficients of the cointegration relations. The break date identified by the Gregory-Hansen methodology is 1993, precisely the last year of the dual exchange rate regime. In view of this, a dummy variable D_t is constructed to account for the regime shift. The dummy is set to unity from 1993 onwards. Hence we estimate the following modified version of equation (11):

$$\ln(E^r) = \mu + \alpha D_t + \beta^H d^H - \beta^F d^F + \eta. \tag{12}$$

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[7] See Baak (2008).

3. Econometric Methodology

Since the main objective is to test for cointegration between real exchange rates and TFP differentials, we first define stochastic cointegration and the deterministic cointegration restriction.[8] Let $y(t)$ be a univariate difference stationary process and $X(t)$ be a multivariate difference stationary process, possibly with nonzero drift. Then $y(t)$ and $X(t)$ may be written as

$$y(t) = \mu_y t + y^0(t), \tag{13}$$

$$X(t) = \mu_X t + X^0(t), \tag{14}$$

where $y^0(t)$ and $X^0(t)$ are difference stationary without drift and μ_y and μ_X are the drift terms.

Definition 1: *$y(t)$ and $X(t)$ are said to be stochastically cointegrated with a normalized cointegrating vector $(1, -\gamma')'$ if there exists a vector γ such that $y^0(t) - \gamma' X^0(t)$ is stationary.*

It is clear from the above definition that stochastic cointegration only requires that the cointegrating vector eliminate the common stochastic trends in the difference stationary variables.

Definition 2: If the normalized cointegrating vector $(1, -\gamma')'$ also satisfies the extra restriction $\mu_y = \gamma' \mu_X$, then it is said to satisfy the deterministic cointegration restriction.

Hence the deterministic cointegration restriction requires that the cointegrating vector, which eliminates the common stochastic trends, also eliminates the deterministic trends in $y(t)$ *and* $X(t)$ arising from the drift terms. Since the key predictions of the Balassa-Samuelson model embodied in equations (6) and (11) imply that the cointegrating vector should remove both the deterministic and stochastic trends, testing for the deterministic cointegration restriction will provide a sharper test of the model than merely testing for stochastic cointegration. In addition, West (1987) and Park (1992) show that imposing the deterministic cointegration restriction on estimators of cointegration vectors also results in efficiency gains.

Since the model implies cointegration, it is desirable to test the null hypothesis of cointegration to control the probability of rejecting a valid economic model. Although estimation methods that have no cointegration as the null are commonly used in the literature, it is well known that these methods have very low power and may fail to reject the null hypothesis with high probability even when the model is actually consistent with the data. By positing cointegration as the null hypothesis, one can control the probability of a Type I error. Park's (1992) canonical cointegrating regression (CCR) procedure is used to test the null hypothesis of stochastic cointegration and the deterministic

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[8]A more detailed discussion is provided by Campbell and Perron (1991) and Ogaki and Park (1997).

cointegration restriction. The CCR estimators are asymptotically efficient and have asymptotic distributions that can essentially be considered as normal distributions, so that their standard errors can be interpreted in the usual way. They also do not require the assumption of a Gaussian VAR structure. Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties than Johansen's (1988, 1991) estimators, even when the Gaussian VAR structure assumed by Johansen is true.[9]

The CCR procedure is applied to the regression

$$y(t) = \mu + \alpha D_t + \sum_{i=1}^{p} \rho_i t^i + \sum_{i=p+1}^{q} \rho_i t^i + \gamma' X(t) + \eta_t, \tag{15}$$

where $y(t)$ is the relative price of nontradables when estimating equation (6) and the real exchange rate when estimating equation (12). Likewise, $X(t)$ is the sectoral TFP differential when estimating equation (6), whereas when estimating equation (12), $X(t) = (d^H, d^F)'$. In order to test for stochastic and deterministic cointegration, superfluous time polynomials up to order q are included in this regression. After the estimated cointegrating vector has been incorporated in the canonical cointegrated regression, if all the coefficients of the added time polynomials are zero, the deterministic cointegration restriction is satisfied. If all coefficients of the added polynomials are zero with the exception of the first-order time polynomial, stochastic cointegration is implied. Let $H(p, q)$ denote the standard Wald statistic to test the hypothesis $\rho_p = \rho_{p+1} = ... = \rho_q = 0$. $H(p, q)$ converges to a χ^2_{q-p} random variable under the null hypothesis of cointegration, and diverges to infinity under the alternative of no cointegration. With appropriate choice of p and q one can test for deterministic and stochastic cointegration. For example, with $p = 0$ and $q = 1$, the $H(0, 1)$ statistic tests the deterministic cointegration restriction. With $p = 1$ and $q = 2$ or 3, the $H(1, q)$ statistics test stochastic cointegration.

4. Data and Empirical Results

4.1. Data

For the U.S., we utilized the STAN industrial database to construct sectoral data on tradable and nontradable output, capital stock and labor hours. Sectoral productivities were constructed as the

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[9] The CCR procedure requires an estimate of nuisance parameters, such as the long-run covariance of the disturbance in the system. The VAR prewhitening method developed by Andrews and Monahan (1992) together with the automatic bandwidth estimator developed by Andrews (1991) are applied to estimate the nuisance parameters. Following the Monte Carlo-based recommendations of Park and Ogaki (1991), the third-step CCR estimates and fourth-step test results are reported. For further details regarding CCR-based estimation and testing, see Ogaki (1993a, b) and Park (1990). All CCR estimation and testing used in this paper are carried out using Ogaki's (1993a) GAUSS-CCR package.

Solow residuals from the Cobb-Douglas production functions for each sector. The following industries were classified as tradable: manufacturing; mining and quarrying; ocean and air transport; wholesale and retail trade; and finance, insurance and business services. The nontradable goods sector comprised: electricity, gas and water; transportation, post and telecommunication; wholesale, retail and catering service; and community, social and personal services. The Chinese GDP classification differs from that of OECD countries in that GDP is classified into the Primary, Secondary and Tertiary sectors. We map these sectors as closely as possible to corresponding U.S. tradable and nontradable sectors. The resulting sample spans from 1980 to 2003. Appendix 1 provides the details of the classification as well as the data sources and methods.

The nominal exchange rate series for the Renminbi against the U.S. dollar was chosen to reflect the market value of the Renminbi as far as possible. Hence from 1980 to 1985, we use the "black market" value of the Renminbi against the dollar[10] rather than the "official" exchange rate. China operated a dual exchange rate system from 1986 to 1993, and foreign and Chinese firms operating in the Special Economic Zones were allowed to negotiate exchange rates among themselves. We use the swap market determined exchange rate for this period[11]. Finally, from 1994 onwards, the dual exchange rate system was abolished and China adopted a unified managed-float regime[12]. We use the market rate for this time period.

There is no available capital stock data for China either at the industry or aggregate levels from official sources. We use two different ways to measure the aggregate capital stock for China. The conventional way to estimating the capital stock is based on cumulating investment data and making an assumption about the (constant) depreciation rate. In a recent work, Holz (2006) argues that the conventional approach is flawed because depreciation is an accounting measure that is not necessarily related to changes in the production capacity of fixed assets. He therefore constructs new estimates of China's aggregate capital stock based on an alternative approach that uses "scrap values" and also adjusts for revaluations of the original values of the fixed assets due to changes in their market prices. Further details of Holz's methodology are provided in Appendix 2. For the purposes of this paper, we take as our leading measure of the aggregate capital stock for China Holz's capital stock series B-C3, which we denote as series "A". Since the measurement of productivity could be sensitive to the choice of the capital stock series, we also construct our own estimate of the aggregate capital stock for China

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[10] The black market value of Renminbi is obtained from the Global Financial Statistics (GFS) database.

[11] As discussed in Huang and Wang (2004), the swap centers formed a platform for a market mechanism outside the central plan and at market rates, and they played a useful role in smoothing the transition from a centrally planned economy to a market economy.

[12] See Huang and Wang (2004) for more details.

by using the conventional approach of cumulating investment data with a constant depreciation rate of 5%. This alternative capital stock series is denoted as series "B".

Once the aggregate capital stock has been estimated, it is then allocated to the tradable and non-tradable sectors in proportion to the share of capital income in that sector. TFPs for the tradable and nontradable sectors are then computed as Solow-residuals from Cobb-Douglas production functions. We present results with estimates of sectoral TFP for both Series A and Series B to assess their robustness to the measurement of capital stocks.

Before testing for cointegration, we pre-test the real exchange rate, the relative prices of nontradables and the sectoral TFP differentials for unit roots. Table 1 reports these results. The first three columns present the results for the ADF test, the Phillips-Perron test, and Park's $J(1,5)$ test for the null hypothesis of difference stationarity against the alternative of trend stationarity. The results are consistent the unit root hypothesis for all variables since none of the test statistics is significant at conventional levels.[13] The last column reports the results of Amsler and Lee's (1995) LM test for testing the null of unit root in the possible presence of a structural break, in view of the 1994 unification of the dual exchange rate system in China. Again, the results are supportive of the unit root hypothesis for all variables except for the U.S. sectoral TFP differential. Since the data are consistent with Assumptions 1 and 2, we proceed next to test for cointegration implied by the Balassa-Samuelson model in equations (6) and (12).

4.2. Results for the Relative Price of Nontradables

Table 2 presents the results of cointegration tests between the relative price of nontradables and the sectoral TFP differential in both China and the U.S. by applying the CCR procedure to equation (6). For China, the $H(1,2)$ and $H(1,3)$ statistics do not reject the null hypothesis of stochastic cointegration at the conventional levels for the TFP estimates constructed based on either series A or B. The deterministic cointegration restriction is also not rejected at the 5% significance level for both series, although it is rejected at the 10% significance level. The point estimates of the coefficient of the sectoral productivity differential are positive and statistically significant for both series. For series A, the estimated coefficient is 0.96, which is strikingly close to the predicted value of unity, and we cannot reject the hypothesis that the estimated coefficient is different from one. For series B, the estimated coefficient is 0.74, which is statistically significantly smaller than the predicted value of unity.

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[13]We also tried the ADF and Phillips-Perron tests without the trend, with almost identical results.

For the U.S., the $H(1,2)$ and $H(1,3)$ statistics do not reject the null hypothesis of stochastic cointegration at the conventional levels of significance. The deterministic cointegration restriction is also not rejected by the $H(0,1)$ statistic at conventional levels. The estimated coefficient of 0.95 is again close to the predicted value of unity and is not significantly different from unity.

Thus, the first key prediction of the Balassa-Samuelson model which states that the sectoral TFP differentials are the key driver of the relative price of nontradables in the long run finds strong support for both China and the U.S. For China, the results for TFP estimates based on series A are slightly better than those for series B.

4.3. Results for the Real Exchange Rate

Table 3 presents the results for cointegration tests between the Chinese-U.S. bilateral real exchange rate and the Chinese and U.S. sectoral TFP differentials. The first column contains the results for TFP estimates based on series A. Both coefficients are correctly signed and significant at the 1% level. Although the point estimates are larger than 1, a two-standard-deviation confidence interval does contain the value 1 as well. The dummy for the change of the exchange rate regime in 1993 is also significant. The $H(1,2)$ and $H(1,3)$ statistics do not reject the null hypothesis of stochastic cointegration at the 5% level, although the $H(1,3)$ statistic is significant at the 10% level. The deterministic cointegration restriction is also not rejected at the conventional significance levels. The second column contains results for TFP estimates based on series B. The results are very similar to those for series A.

We also plot the actual real exchange rates and their estimated "equilibrium" values based on the Balassa-Samuelson model. Along with the estimated equilibrium values, we also plot their 95% confidence bands which are helpful in assessing statistically significant departures from the long run equilibrium values. Panel A of Figure 1 shows the results for the real exchange rate using series A. It shows that the model can capture the depreciating trend of the real exchange rate from 1980 to 1990 and its subsequent appreciation from 1991 through 1996. The real exchange rate moves out of the 95% confidence bands during 1984-87 and 1989-90. The Chinese Yuan appears to be significantly overvalued during 1984-87, which is consistent with the conventional wisdom of economic observers of China (see, e.g., Chow (2002)). The overvaluation is generally attributed to high growth rates of money supply exceeding 30% per year and the consequent high inflation rates during this period. By contrast, in 1989-91, the exchange rate appears to be significantly undervalued. This effect might plausibly be related to the 1989 Tiananmen incident, following which there was a sharp temporary

decline in foreign investment and trade with China. In the most recent decade from 1994 to 2003, however, there is no indication of a statistically significant misalignment of the real exchange rate. Although the real exchange rate has a tendency to be slightly undervalued, such deviations are small (about 5 percent in 2003) and tend to dissipate over time. In fact, the actual and predicted real exchange rates are almost identical in 2003, which is our last observation. This analysis suggests that recent claims about the drastic undervaluation of the Renminbi are quite plausibly exaggerated. Panel B of Figure 1 shows the actual and predicted values of the real exchange rate for series B, which are qualitatively very similar as those for series A.

To sum up, regardless of which capital stock series is used, the second key prediction of the Balassa-Samuelson model is also supported by the evidence. Changes in the Chinese and U.S. sectoral TFP differentials seem to explain most of the long term movements in the Chinese real exchange rate.

5. Exploring the Role of Net Foreign Assets

As mentioned in the introduction, in addition to changes in sectoral TFP differentials, accumulation of foreign assets or debt may also have a significant impact on the long run real exchange rate movements of a country. Theoretical work by Hooper and Morton (1982), Gavin (1991) and more recently by Lane and Milesi-Ferretti (2002) and Thoenissen (2005) suggests that rising net foreign assets should in the long run be associated with a real appreciation of the real exchange rate. Empirical work by Égert, Lommatzsch and Lahrèche-Révil (2006), Gagnon (1996) and Lane and Milesi-Ferretti (2004) generally corroborates this positive correlation. We use the data provided in Lane and Milesi-Ferreti's (2006) seminal work as our measure of the net foreign asset position of China. Similar to Gagnon (1996), net foreign assets are scaled by the volume of trade since in most theoretical models, equilibrium is achieved via the impact of the real exchange rate on the trade sector of the economy.

Table 4 presents the results for cointegration tests by adding the net foreign asset variable to the model in addition to the sectoral TFP differentials. For series A, the coefficient of the net foreign asset variable is positive and statistically significant. However, for series B the coefficient of the net foreign asset variable is not significant. For both series A and B, the coefficients of the sectoral TFP variables retain their correct signs and statistical significance. The $H(1,2)$ and $H(1,3)$ tests do not reject the null hypothesis of stochastic cointegration for either series A or B. The deterministic cointegration restriction is also not rejected for either series by the $H(0,1)$ test statistic. Indeed the p-values of the $H(0,1)$, $H(1,2)$, and $H(1,3)$ test statistics are noticeably larger, for both series A and

B, compared with the same statistics when net foreign assets are excluded. Hence the inclusion of the net foreign asset variable appears to enhance the cointegration results.

Panels A and B of Figure 2 plot the estimated equilibrium real exchange rate implied by the model along with 95% confidence bands for series A and B, respectively. The graphs are qualitatively similar to Figure 1, which excludes net foreign assets, in that the real exchange rate appears significantly overvalued in 1985-87 and significantly undervalued in 1989-91. The fit between the actual real exchange rate and its estimated equilibrium value is even closer, relative to Figure 1, in the most recent decade from 1994 to 2003. Hence the result that the real exchange rate is close to its equilibrium value in recent years is robust to the inclusion of net foreign assets.

6. Conclusions

Using a carefully constructed TFP measure of productivity, our findings generally confirm the validity of the Balassa-Samuelson effect in determining the long-run movements of the Chinese real exchange rate against the U.S. dollar. We find that changes in the sectoral TFP differential can explain the evolution of the relative price of nontradable to tradable goods within both China and the U.S. as well as the bilateral Renminbi-Dollar real exchange rate. It is noteworthy that the model identifies 1984-87 as a period of significant overvaluation and 1989-91 as a period of significant undervaluation, which is consistent with received wisdom about China's exchange rate from economic observers.

When we compare the actual real exchange rates with the equilibrium values derived from the productivity-based model in more recent years, we find that while the Renminbi has been somewhat undervalued against the dollar, the misalignment is relatively small and statistically insignificant. However, given the significant uncertainties involved in estimating equilibrium real exchange rate with China's data, as well as the plethora of possible approaches that could potentially be employed, these results should be interpreted with caution.

Appendix 1: Data Description

1.1 Measurement of Factor Inputs

The labor input is taken as the total number of employed workers multiplied by the estimated number of work hours. As no official organization or survey reports the actual number of work hours, we follow Hu and Khan (1997) and use 48 hours per week as an estimate since the standard work hours mandated by the Chinese State Council is 48 hours per week.[14] In the official data source, the labor force is exhaustively divided into the primary, secondary and tertiary sectors. The secondary sector is further divided into the industry and construction components. In this paper, the labor force in the primary and manufacturing industries are included in the tradable sector, and the rest are included in the nontradable sector.[15]

Labor compensation is taken as the total number of employed workers by industry multiplied by the average annual wage rate.[16]

The capital input (series B) is measured using the perpetual inventory approach, similar to that used in Chow (1993), Chow and Li (2002), Li (2002) and Feenstra and Kee (2004). This method depends on the investment series, which is cumulated and scrapped based on the depreciation assumption. Following Kim and Lau (1995), the initial capital stock is taken to be five times the real gross fixed capital formation in the base year, which is set to 1970.[17] The fixed capital stock is then approximated by adding to the initial stock the real gross fixed capital formation since the base year, which includes the non-residential building, other construction, and machinery and equipment. The net real investment accumulation is treated as a net increase in capital stock. Mathematically,

$$K_t = (1 - \delta) K_{t-1} + RGI_t$$

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[14]Holz (2006) provides a detailed estimate of the work hours for different industries in five selected years (1995, 2001-2004). Holz's estimate is 44.9-45.5 hours per week during 2001-2004, which is quite close to our proxy.

[15]Even though the labor force data at the industry level is available up to 2002 (see Appendix 15 on p.240 of Holz (2006)), we do not use the industry breakdown provided in this sectoral report form as the sum of the industries does not add up to the total labor force and nobody knows into which industries the residual belongs.

[16]For the agricultural industry, the total labor payroll is computed as the share of the labor compensation in the total gross output as given in the China Statistical Yearbook (since the early 1990s) and "Zhong Guo Guo Nei Sheng Chan Zong Zhi He Suan" (since 1978).

A possible alternative way to calculate the labor compensation in the other industries is by dividing the labor remuneration in the national income accounts with the number of laborers in the industries. However, since the income approach data to the calculation of GDP are highly unreliable (with major revision and statistical break in 2005), this computation approach is not used in this study.

A detailed summary of the labor remuneration in 1978-2002 is provided in Holz (2006) p.185, 249, 257-259.

[17]Real gross fixed capital formation = Nominal gross fixed capital formation / GDP deflator for the secondary industries.

where RGI_t is the real gross investment during period t, and $\delta = 0.05$ is the depreciation rate.

The deflators of the tradable and nontradable sectors are calculated as follows: the producer price index[18] serves as the proxy of the tradable sector, while the deflator of the nontradable services sector is the geometric difference between the aggregate GDP deflator and the tradable sector's deflator.[19]

$$P = (P_T)^\alpha (P_N)^{1-\alpha}$$
$$P_N = (P/(P_T)^\alpha)^{1/(1-\alpha)}$$

where P and P_T are the GDP deflator[20] and the producer price index respectively. α is share of the tradable sector in the total GDP.

In China, the classification of GDP by sector is different from the classification used in other countries. The GDP in China is classified into the primary, secondary and tertiary industries. The tertiary industry is subdivided into 1) transportation, post and telecommunications, 2) wholesale, retail trade and catering service and 3) others. We take the output of the primary industry as the proxy of the output for the agriculture industry. The secondary industry is subdivided into 1) industry (which includes mining and quarrying, manufacturing and utilities) and 2) construction. The output for the sectors "finance, insurance and business services" and "community, social and personal services" from 1990 to 2003 are taken from the output values measured via the value-added approach (Table 3-5 of the China Statistical Yearbook entitled "Value-Added of the Tertiary Industry"). As the data from 1980 to 1989 is not available, we use the changes of the category "others" of the tertiary industry as a proxy for the changes in these two industries.

The cumulative capital stock is then distributed to the tradable and nontradable sectors according to the share of capital income in the sector. The labor share is estimated as the labor compensation of that sector divided by the total sectoral output. The capital share is then taken as one minus the labor share of that sector. Capital income in each sector is the product of the total output of each sector and the share of capital income. The capital stock of each individual industry is the product of the

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[18]Producer Price Index reflects the trend and degree of changes in the general ex-factory prices of all industrial products during a given period, including sales of industrial products by an industrial enterprise to all units outside the enterprise, as well as sales of consumer goods to residents. It can be used to analyze the impact of ex-factory prices on gross output value and value-added of the industrial sector.

[19]Even though this method may be a bit indirect, we make use of this procedure because the official data on sectoral price deflators is highly problematic.

[20]Note that the implicit GDP deflator of the years 1993-2004 (and minimally in 1978-1992) changed following the 2004 economic census. Here we use the unrevised earlier implicit GDP deflator as it is more relevant to the market perception at the time of exchange rate determination. Similar view is shared by Holz (2006) as well as Molodtsova, Nikolsko-Rzhevskyy and Papell (2007).

national capital stock and the ratio of capital income of each industry to the total capital income.[21]

1.2 Data Sources

The sample period is 1980-2003. The data is primarily collected from various issues of the Statistical Yearbook of China, and supplemented by the CEIC database. The detailed sources are as follows:

- GDP at current price: China Statistical Yearbook Table 3-1 ("Gross Domestic Product").

- GDP at constant price: CEIC Table CN.A03: "GDP by Industry (Index: Previous Year=100)".

- Gross fixed capital formation: China Statistical Yearbook Table 3-14 ("Structure of Gross Domestic Product by Expenditure Approach")

- Employment by industry: China Statistical Yearbook Table 5-6 ("Number of Employed Persons at the Year-end by Sector")

- Average salary for workers: China Statistical Yearbook Table 5-24 ("Average Wage of Staff and Workers by Sector")

- Producer price index: CEIC Table CN.I13: ("Producer Price Index: Industrial Products")

- Imports and Exports: China Statistical Yearbook Table 17-1 ("Foreign Trade and Economic Cooperation")

- CPI: CEIC Table CN.I20 and Table US.I01 (Consumer Price Index: Urban Area)

- Net foreign asset: International Financial Statistics (IFS)

- Swap exchange rate: provided by the State Administration of Foreign Exchange (SAFE)

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[21] The same approach is used in Li (2002) to derive the capital stock by industry.

Appendix 2: Holz's (2006) Capital Stock Series

The estimate of the aggregate capital stock series is an important part of productivity studies. Recently, Holz (2006) has made a serious effort to provide new capital estimates for China based on the logic of the accounting system. His work has inspired meaningful discussion among researchers in this area.[22] Holz makes several modifications to the conventional approach in order to improve the measurement of capital stock. In particular, he points out two shortcomings of the conventional approach. First, he argues that applying a constant depreciation rate to calculate changes in the capital stock is unreasonable as depreciation is an accounting measure that bears no necessary relation to changes in the production capacity of fixed assets. Mechanically applying a constant depreciation rate to the capital stock is inappropriate as the machines are likely to potentially operate at the same capacity as at its purchasing date until the time when the machines are decommissioned. As such, he provides estimates of the scrap values rather than following the standard way of applying a constant depreciation rate. Second, he argues that the use of investment expenditure is incorrect because not all expenditures lead to increases in fixed assets. The difference could be due to waste or time lags between the time when money is being spent and when the completed fixed assets are ready for use.

In view of this, Holz suggests an alternative method to construct capital stock estimates. This new method is based on the following identity that specifies the changes of the original value of fixed assets over time:

$$OFA_t - OFA_{t-1} = \text{investment}_t - \text{scrap value}_t + \text{revaluation}_t$$

In this identity, "OFA" denotes the original value of fixed assets, "investment" refers to the effective investment which is defined as the newly increased fixed assets through investment rather than to investment expenditure, "scrap value" refers to the original value of decommissioned fixed values, and "$revaluation$" refers to the revaluations of the original values of the fixed assets to current market prices.

Holz reports several different estimated capital stock series constructed based on different ways to calculate the effective investment. Holz's B-C3 series derives the effective investment in the non-state-owned units (non-SOUs) based on the real growth rate of non-SOU industrial gross output value. Holz's B-C4 series uses the annual non-SOU investment (gross fixed capital formation less SOU

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[22]See, for example, Chow (2006).

investment) and turns it into effective investment using the non-SOU transfer rate.[23]   The results reported in this paper are based on the B-C3 series. The results for the B-C4 series are qualitatively similar and can be made available by the authors upon request.

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[23] Transfer rate is defined as the ratio of effective investment to investment expenditure. Its value is explicitly provided in the official statistics.

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Table 1: Unit Root Tests

	ADF (with trend)	Phillips-Perron (with trend)	J(1,5) (with trend)	LM
	H_0 : unit root			
RER	-1.9654 (0.5871)	-1.5363 (0.7863)	5.5550	-2.6682
Q_{China}	-2.2019 (0.4665)	-2.2019 (0.4665)	2.5701	-2.2895
d_{China}	-1.6614 (0.7354)	-1.6614 (0.7354)	2.4120	-2.5521
Q_{US}	-2.6353 (0.2697)	-1.7645 (0.6885)	2.7707	-2.8226
d_{US}	-1.9807 (0.5806)	-2.1115 (0.5129)	1.2483	-3.2580*

Notes: P-values are in parentheses. * denotes significance at the 5% level.

The critical values of the LM test are given in Schmidt and Phillips (1992).

Table 2

Cointegration Tests Between the Relative Price of Nontradables and the Sectoral TFP Differential

Variable	China (Series A, H-B3)		China (Series B)		US	
	Coefficients	Standard Error	Coefficients	Standard Error	Coefficients	Standard Error
Constant	-0.0590***	(0.0205)	0.0494*	(0.0285)	0.3755***	(0.0111)
d	0.9619***	(0.0404)	0.7368***	(0.0551)	0.9474***	(0.0279)
Tests of deterministic and stochastic cointegration:						
$H(0,1)^b$	2.8547*	(0.0911)	3.3857*	(0.0658)	2.0473	(0.1525)
$H(1,2)^b$	0.9359	(0.3333)	1.0340	(0.3092)	0.4593	(0.4980)
$H(1,3)^b$	1.4988	(0.4726)	1.6544	(0.4373)	2.0630	(0.3565)

Notes: Standard errors are in parenthesis. ^{b}H(0,1) tests the null hypothesis of deterministic cointegration restriction, whereas H(1,2) and H(1,3) test the null hypothesis of stochastic cointegration. P-values are in parentheses for H(0,1), H(1,2) and H(1,3).

*, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels respectively.

Table 3: Cointegration Tests Between the Chinese Real Exchange Rate and the Sectoral TFP Differentials of China and the U.S.

	Series A (H-B3)		Series B	
	Coefficient	Standard Error	Coefficient	Standard Error
Constant	-3.9212***	(0.3082)	-3.8372***	(0.2816)
d_{China}	1.5242***	(0.2967)	1.4617***	(0.2534)
d_{US}	-1.6226***	(0.5018)	-1.4897***	(0.4708)
Dummy	1.1482***	(0.2495)	1.1473***	(0.2337)
	Tests of Deterministic and Stochastic Cointegration			
	Statistic	p-value	Statistic	p-value
$H(0,1)^b$	1.1602	(0.2814)	1.3205	(0.2505)
$H(1,2)^b$	1.3450	(0.2462)	1.4625	(0.2265)
$H(1,3)^b$	5.7015*	(0.0578)	5.4590*	(0.0653)

Notes: Standard errors are in parentheses. ^{b}H(0,1) tests the null hypothesis of deterministic cointegration restriction, whereas H(1,2) and H(1,3) test the null hypothesis of stochastic cointegration. P-values are in parentheses for H(0,1), H(1,2) and H(1,3).

*, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels respectively.

Table 4: Cointegration Tests With Net Foreign Assets

| Variable	Series A		Series B	
	Coefficient	Standard Error	Coefficient	Standard Error
Constant	-3.5392***	(0.4113)	-3.8797***	(0.3350)
d_{China}	0.9697**	(0.3734)	1.4231***	(0.2508)
d_{US}	-1.4061***	(0.4697)	-1.6148***	(0.4387)
NFA/Trade	0.6148*	(0.3185)	0.3366	(0.2684)
Dummy	1.0690***	(0.2175)	1.2297***	(0.2228)
Tests of Deterministic and Stochastic Cointegration				
	Statistic	p-value	Statistic	p-value
H(0,1)	4.9950e-006	(0.9982)	0.0169	(0.8966)
H(1,2)	0.0687	(0.7932)	0.0241	(0.8766)
H(1,3)	0.3190	(0.8526)	1.3719	(0.5036)

Notes: Standard errors are in parentheses. H(0,1) tests the null hypothesis of deterministic cointegration restriction, whereas H(1,2) and H(1,3) test the null hypothesis of stochastic cointegration. P-values are in the parentheses for H(0,1), H(1,2) and H(1,3).

*, ** and *** denote significance at 10 percent, 5 percent and 1 percent levels respectively.

Figure 1A: Actual and Equilibrium Real Exchange Rates in China, along with the 95% Confidence Interval (Productivity Model with Series A)
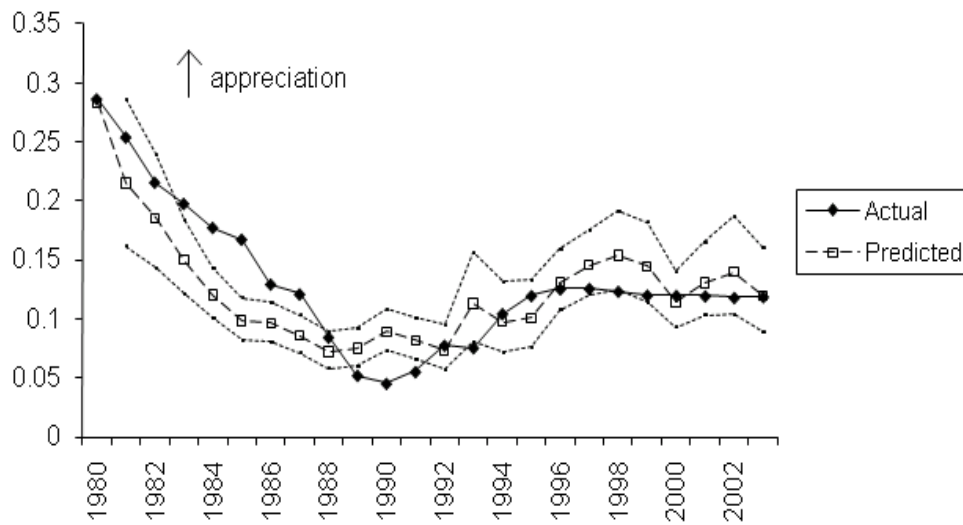
Figure 1B: Actual and Equilibrium Real Exchange Rates in China, along with the 95% Confidence Interval (Productivity Model with Series B)
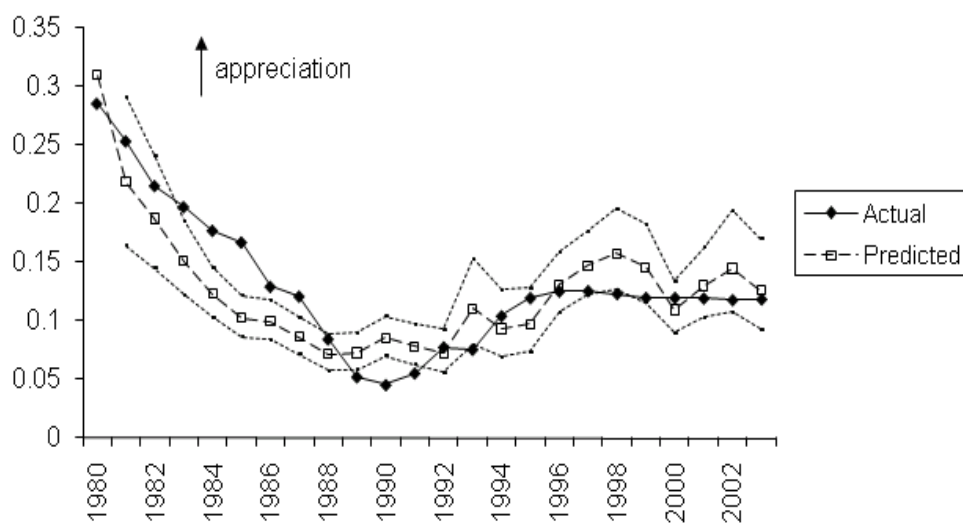
Figure 2A: Equilibrium and Actual Real Exchange Rates in China, Augmented by NFA/Trade (Series A)
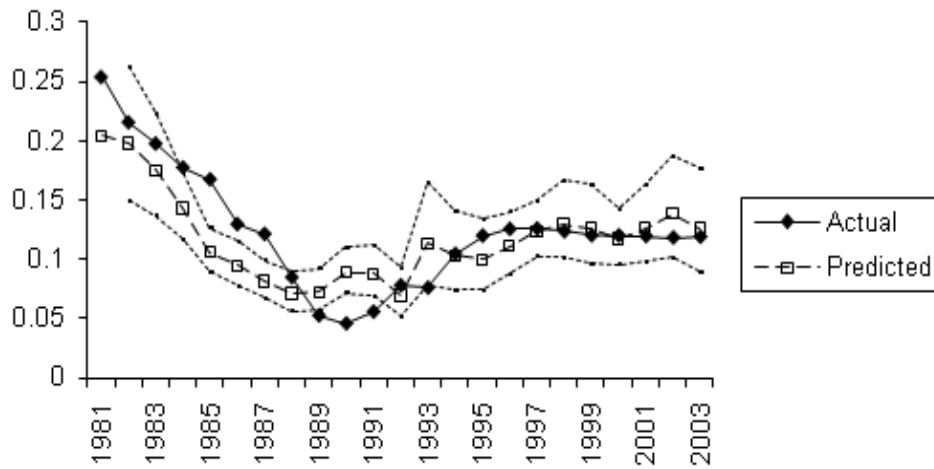
Figure 2B: Equilibrium and Actual Real Exchange Rates in China, Augmented by NFA/Trade (Series B)