University of Leicester

DEPARTMENT OF ECONOMICS

A TEST OF THE BALASSA-SAMUELSON EFFECT APPLIED TO CHINESE REGIONAL DATA

Qian Guo, Loughborough University, UK
Stephen G Hall, University of Leicester, UK

Working Paper No. 08/8
March 2008

A Test of the Balassa-Samuelson Effect

Applied to Chinese Regional Data

By

Qian Guo[1]

And

Stephen G Hall[2]

1 Business School, Loughborough University, Leicester LE11 3TU, United Kingdom
Tel: 44 (0)1509 22 3177 (office) E-mail: q.guo@lboro.ac.uk

2 Department of Economics University of Leicester and National Institute of Economic and Social Research, London

Department of Economics, University of Leicester, University Road, Leicester LE1 7RH, United Kingdom Tel: 44 (0) 116 252 2827 Fax: 44 (0)116 252 2908 E-mail: s.g.hall@le.ac.uk

National Institute of Economic and Social Research, 2 Dean Trench Street, Smith Square, London SW1P 3HE, United Kingdom Tel: 44 (0) 20 7654 1929 Fax: 44 (0) 20 7654 1900 E-mail: shall@niesr.ac.uk

A Test of the Balassa-Samuelson Effect Applied to Chinese Regional Data

ABSTRACT

In this paper we investigate the relevance of the Balassa-Samuelson effect to the determination of regional inflation in China, for the period 1985 – 2000. To do this, we first construct annual measures of Chinese inflation and industry input on regional and sectoral basis. Then we generalize the Asea and Mendoza (1994) settings to consider asymmetric productivity shocks across sectors. Testing this model on Chinese Regional Data aid of non-stationary panel data techniques, it shows that our extended theoretical model is a good empirical representation of the Chinese data which supports the Balassa-Samuelson effect. Moreover, we are able to test the Asea and Mendoza (1994) version of our general model and find that the restrictions are rejected.

1. INTRODUCTION

The main proposition of the Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964) is that high productivity growth of the tradable sector compared to the non-tradable sector leads to a rise in the relative price of nontradables, which puts upward pressure on a country's real exchange rate. Previous empirical analysis of the Balassa-Samuelson effect can be broadly broken into two groups. The first group of literature is static in nature and examines the major components of theory straightforwardly (see, in particular, Canzoneri, Cumby, and Diba 1999; Chinn 2000; Drine and Rault 2002, 2003; Marston 1990; Micossi and Milesi-Ferretti 1994; Strauss 1995; Vikas and Ogaki 1999). The second group is concerned with intertemporal equilibrium model incorporating nontradables, which, in general, specifies production and consumption in the context of intertemporal optimization. In such a strand of literature, Chinn (1995) and (1996), De Gregorio and Wolf (1994), Froot and Rogoff (1991), Obstfeld (1993), and Rogoff (1992) adopt the unbalanced growth framework and examine the time series implication of the Balassa-Samuelson effect. On the other hand, Asea and Mendoza (1994) propose a model with balanced growth features and are concerned with the cross-sectional impact of the theory. Our analysis extends the Asea and Mendoza (1994) approach and differs from them in that we model the two-country and two-sector world where shocks to technologies are heterogeneous across sectors.

For the empirical analysis, we turn to a Regional Data set for China which has been especially constructed for this work. The reason why estimation based on regional data is interesting for China lies in the fact that the inflation and productivity trends across Chinese regions have varied enormously.[1] Such diversity would make it rather difficult to apply time series methods to aggregated data. The central contribution of our study is the construction of a large disaggregate database on China. And so the details of the data and the construction of the empirical counterparts to the theoretical variables merit some discussion (additional details are in Part 3). China is composed of twenty-two provinces (Anhui, Fujian, Gansu, Guangdong, Guizhou, Hainan, Hebei, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Jiangxi, Jilin, Liaoning, Qinghai, Shanxi, Shaanxi, Shandong, Sichuan, Yunnan, and Zhejiang), five autonomous regions (Guangxi, Inner Mongolia, Ningxia, Xinjiang, and Tibet), four municipalities (Beijing, Chongqing,[2] Shanghai, and Tianjin), and two special administrative regions (Hong Kong and Macau). Hong Kong and Macau are not within our scope of study due to different political and administrative systems compared to mainland China. The data for Chongqing have been integrated with those for Sichuan due to the lack of data before 1997. Thus, our sample consists of thirty regions in total. The data are yearly in frequency, for the period 1985 – 2000. The real exchange rate is defined as the bilateral real exchange rate between each region of China and the United States, adjusted to the difference in the GDP deflators of each Chinese region and the United States. The prices of tradables and nontradables are proxied by the GDP deflator for the Chinese agriculture and industry, and for "other," which includes construction, transportation, storage, postal and telecommunications services, wholesale, retail trade, and catering services.[3] These data are calculated as the ratio of the nominal to real

---

[1] For example, from 1992 to 1999, the average annual rates of variation for consumer price index (CPI) have ranged from 8.1% in Hainan to 11.5% in Beijing. Over the decade of the 1990s, total inflation in Beijing exceeds that in Hainan by 40% (Guillaumont Jeanneney and Hua, 2002).

[2] Chongqing was formerly (until 14 March 1997) a sub-provincial city within Sichuan province.

[3] The tradable and nontradable categorization is on the basis of De Gregorio and Wolf (1994) criterion, which is discussed in more detail in Part 3.

GDP index, both at 2000 constant price for each sector and region. Attempting a procedure for calculating sectoral productivity is a difficult enterprise. Previous analyses of productivity-based models of the real exchange rate have employed labour productivity,[4] which is calculated as GDP per hour worked, or, the total factor productivity (TFP)[5] computed as the Solow residuals. As De Gregorio, Giovannini, and Krueger (1994) show, the use of labour productivity might be tainted by the demand effects, such as labour shedding, or, the unadjustment process for part-time workers especially in the agricultural sector (Chinn 2000). In this study, we use the TFP approach constructed from the real GDP, capital stock, employment,[6] and factor returns for each sector.[7] The output, employment, and wage rate necessary to the construction of factor returns are drawn from the China Statistical Yearbooks (CSYB). The capital stocks are approximated through investment data, which are also from CSYB.

In this paper, we test the Asea and Mendoza (1994) version of our general model on the Chinese Regional Data, using recently developed non-stationary panel data methods. The results show that eq. (I) of our extended model is a reasonable empirical representation of the Chinese Balassa-Samuelson effect. Thus, we are able

[4] See, for example, Bergstrand (1991), Canzoneri, Cumby, and Diba (1999), Froot and Rogoff (1991), Hsieh (1982), and Marston (1987).

[5] See, for example, Chinn (1995), Chinn and Johnston (1996), De Gregorio, Giovannini, and Krueger (1994), and De Gregorio, Giovannini, and Wolf (1994).

[6] Due to the lack of data for labour hours, we follow most studies and use the total employment data as a proxy.

[7] However, interpreting the change in TFP as exogenous supply shocks is problematic. (Evans 1992) shows that measured Solow residuals are Granger caused by money, interest rate, and government expenditure. Also, the reliability of the TFP is likely to be affected by mis-estimates of the capital stock and labour shares.

to test the Asea and Mendoza (1994) model restrictions relative to our more general model and these restrictions are rejected.

Our study is organized as follows. Part 1 is the introductory section. Part 2 describes the general equilibrium model that motivates our empirical tests. Part 3 presents the data and variable constructions. Part 4 summarizes our empirical evidence. Part 5 concludes.

2. THE BALASSA-SAMUELSON EFFECT: A GENERAL EQUILIBRIUM APPROACH

The main theoretical framework on which we base our empirical work is the two-country and two-sector general equilibrium analysis first proposed by Asea and Mendoza (1994). The original work focuses on the long-run balanced growth, assuming that the productivity shocks follow transitory deviations from the steady-state growth path.[8] We argue that the productivity shifters, and so the TFPs, are heterogeneous across sectors in a real world situation. In what follows, we first briefly review the Balassa-Samuelson framework. Then we extend the Asea and Mendoza (1994) settings to consider a more general assumption on technologies.

2A. The Balassa-Samuelson hypothesis

In the Balassa-Samuelson hypothesis, high growth is made possible by high productivity growth, with differential sectoral growth rates that cause the inflation differentials among sectors. The relative price of nontradables is expected to rise faster in countries with faster growth, since the differential in inflation rates have to

[8] As a result, the sectoral random disturbances to technologies cancel each other out in the closed-form solutions of the relative price.

widen to make the overall growth rate higher (Ito, Isard, and Symansky 1997). In a schematic way, there are three major components of the Balassa-Samuelson theory. The first is that a faster increase of tradable sector productivity than of nontradable ones leads to an increase in relative price of nontradables. The second is that the Purchasing Power Parity (PPP) holds for tradable goods. The third component, which is based on the previous two assumptions, says that high productivity growth of the tradable sector comparing to the nontradable sector causes a rise in the relative price of nontradables, which puts upward pressure on a country's real exchange rate.

2B. The Firms

Suppose that there are two industries in the economy, each containing a large number of homogeneous firms, producing tradable (T) and nontradable (N) goods subject to the constant-returns-to-scale production function with the labour-augmenting [9] technological progress: $Y_t^i = A_t^i (K_t^i)^{1-\alpha_i} (X_t N_t^i)^{\alpha_i}$, $i = T, N$, where Y_t^i is the output value, A_t^i is a random disturbance to TFP, K_t^i is the predetermined value of the capital stock, X_t is an index of technology, N_t^i is the flow of labour hours, and α^i is the labour share. The Solow residual, or, the TFP, for each sector i is: $\theta_t^i = A_t^i (X_t)^{\alpha_i}$, $i = T, N$. The production function has standard neoclassical properties. It is concave and twice continuously differentiable, satisfies the Inada conditions and implies that both factors of production are essential.

Let us denote the flow of leisure hours as L. The domestic labour market is in

[9] For the economy to grow at a constant rate, technological progress must take the labour-augmenting form (Solow 1956).

equilibrium ex-ante where $1-L=N^T+N^N$. Labour is internationally immobile but can migrate instantaneously between sectors within the economy. There is, however, no economy-wide resource constraint for capital comparable to the labour constraint. The evolution of capital is: $K_{t+1}=(1-\delta)K_t+I_t$, where the investment I_t is the amount of current output stored for use in the production in the next period; δ is the rate of deprecation.

The assumption of perfect mobility of labour across sectors ensures that the wage rates are identical in two sectors in the long-run, which shows: $p_t^N=\frac{\alpha_T \frac{Y_t^T}{N_t^T}}{\alpha_N \frac{Y_t^N}{N_t^N}}$, where p_t^N is the relative price of nontradable goods, i.e. $p_t^N=\frac{P_t^N}{P_t^T}$. Substituting the production functions into p_t^N, we have:

$$p_t^N=\frac{\alpha_T(\theta_t^T)^{\frac{1}{\alpha_T}}(\frac{K_t^T}{Y_t^T})^{\frac{1-\alpha_T}{\alpha_T}}}{\alpha_N(\theta_t^N)^{\frac{1}{\alpha_N}}(\frac{K_t^N}{Y_t^N})^{\frac{1-\alpha_N}{\alpha_N}}}\ . \qquad (1)$$

The relative price expression (1) we obtain is an extended version of eq. (27) in Asea and Mendoza (1994).

2C. The Households

We assume that the economy consists of infinitely lived consumers, who maximize their discounted sum of the expected utility:

$$\underset{C_t^T,C_t^N,L_t}{Max}\ E[\sum_{t=0}^{\infty}D^tU(C_t^T,C_t^N,L_t)],$$

where E is the expectation operator; $D=\frac{1}{1+\rho}$ is the discount factor, 0<D<1, ρ is the subjective discount rate; $U(C_t^T, C_t^N, L_t)$ is the instantaneous utility;[10] C_t^T and C_t^N are the consumption expenditures on tradable and nontradable goods, respectively, subject to the budget constraint:

$$C_t^T + P_t^N C_t^N$$

$$=(r_t^T K_t^H + r_t^{T*} K_t^F + P_t^N r_t^N K_t^N)+(w_t^T N_t^T + P_t^N w_t^N N_t^N)$$

$$-\gamma(K_{t+1}^H + K_{t+1}^F + P_t^N K_{t+1}^N) + (1-\delta)(K_t^H + K_t^F + P_t^N K_t^N) - \gamma R_t b_{t+1} + b_t \,,$$ [11]

where we take the tradable goods as the numeraire, with a common price of one in each of two countries, home (H) and foreign (F); the nontradable goods have distinct

---

[10] The instantaneous utility function $U(C_t^T, C_t^N, L_t)$ has the form of the constant-intertemporal-elasticity-of-substitution, where the inverse of the elasticity of marginal utility is constant. It is assumed log linear in its two arguments: $U(C_t^T, C_t^N, L_t) = \ln[U(C_t^T, C_t^N)L_t^{\omega}]$, where ω is the elasticity of leisure. The utility function $U(C_t^T, C_t^N)$ has the form of constant-elasticity-of-substitution: $U(C_t^T, C_t^N) = [\Omega(C_t^T)^{-\mu} + (1-\Omega)(C_t^N)^{-\mu}]^{-\frac{1}{\mu}}$, where μ>1, μ≠0, $El_{C_t^T, C_t^N} = \frac{1}{1+\mu}$; Ω is the share of the composite consumption and 0<Ω<1. The specifications of the utility functions $U(C_t^T, C_t^N, L_t)$ and $U(C_t^T, C_t^N)$ allow us to obtain some specific form of the instantaneous utility such that $U(C_t^T, C_t^N, L_t) = \frac{\{[\Omega(C_t^T)^{-\mu} + (1-\Omega)(C_t^N)^{-\mu}]^{-\frac{1}{\mu}} L_t^{\omega}\}^{1-\sigma}}{1-\sigma}$, where σ >0 is the inverse of the elasticity of the intertemporal substitution.

[11] In the absence of the government sector, the representative households' consumption expenditures are financed by the value of total output minus investment plus net foreign assets. In addition, the assumption of perfect international capital mobility means that resources can always be borrowed abroad and turned into domestic capital. Thus, the total real returns paid to the households include the ones on the capital stock in the foreign tradable sector.

home and foreign prices; an asterisk denotes the foreign country; r_t^i is the real interest rate; w_t^i is the real wage rate; γ is the nominal interest rate; R_t is the inverse of the real gross rate of return paid on international bonds; b_t is the net foreign assets accumulated by the households.

2D. Competitive Equilibrium and Relative Price

The Lagrangean maximization problem with respect to C_t^T and C_{t+1}^T yields:

$$\frac{U_1(t)}{DE[U_1(t+1)]} = \frac{\lambda_t}{\lambda_{t+1}}, \tag{2}$$

where we denote $U'(C_t^T, C_t^N, L_t)$ as U1(t), $U'(C_{t+1}^T, C_{t+1}^N, L_{t+1})$ as U1(t+1), $U'(\cdot)$ is the differentiation operator. Another set of market clearing conditions are obtained through differentiating the Lagrangean with respect to $b_{t+1,} K_{t+1}^H, K_{t+1}^F, K_{t+1}^N$, respectively:

$$\lambda_t \gamma Rt = \lambda_{t+1} \tag{3}$$

$$\lambda_{t+1}[r_{t+1}^T + (1-\delta)] = \lambda_t \gamma \tag{4}$$

$$\lambda_{t+1}[r_{t+1}^T * + (1-\delta)] = \lambda_t \gamma \tag{5}$$

$$\lambda_{t+1} P_{t+1}^N [r_{t+1}^N + (1-\delta)] = \lambda_t P_t^N \gamma \tag{6}$$

By substituting eq. (3) - (6) into (2), one can see the following equilibrium conditions:

$$\gamma R_t U_1(t) = DE[U_1(t+1)]; \tag{7}$$

$$\gamma U_1(t) = DE[U_1(t+1)][r_{t+1}^T + (1-\delta)]; \tag{8}$$

$$\gamma U_1(t) = DE[U_1(t+1)][r_{t+1}^T * + (1-\delta)]; \tag{9}$$

$$P_t^N \gamma U_1(t) = DE[U_1(t+1)] P_{t+1}^N [r_{t+1}^N + (1-\delta)]. \tag{10}$$

From eq. (8) and (10), it follows that in a deterministic stationary equilibrium with perfect sectoral capital mobility, the marginal product of capital in the tradable and nontradable sectors are equalized (Asea and Mendoza 1994), which is:

$(1-\alpha_T)\frac{Y_t^T}{K_t^T} = (1-\alpha_N)\frac{Y_t^N}{K_t^N}$. If we incorporate this equilibrium condition into our relative price expression (1), we further generate:

$$p_t^N = \frac{\alpha_T (\theta_t^T)^{\frac{1}{\alpha_T}}}{\alpha_N (\theta_t^N)^{\frac{1}{\alpha_N}}} (\frac{1-\alpha_N}{1-\alpha_T})^{\frac{\alpha_N - 1}{\alpha_N}} (\frac{K_t^T}{Y_t^T})^{\frac{1-\alpha_T}{\alpha_T} - \frac{1-\alpha_N}{\alpha_N}} . \qquad (11)$$

Eq. (11) may be re-written in terms of the steady-state investment as:

$$p_t^N = \frac{\alpha_T (\theta_t^T)^{\frac{1}{\alpha_T}}}{\alpha_N (\theta_t^N)^{\frac{1}{\alpha_N}}} (\frac{1-\alpha_N}{1-\alpha_T})^{\frac{\alpha_N - 1}{\alpha_N}} (\frac{\frac{I_t^T}{Y_t^T}}{\gamma - (1-\delta)})^{\frac{1-\alpha_T}{\alpha_T} - \frac{1-\alpha_N}{\alpha_N}} . \qquad (12)$$

The expressions (11) and (12) we obtain are extended versions of the relative price equations (see Asea and Mendoza 1994: p.251-252 for comparison) due to our generalized assumptions on productivity shifters.

2E. The long-run real exchange rate

Suppose that the representative households have the constrained budget minimization problem for unit utility: $\underset{C_t^T, C_t^N}{Min}\ UC = P_t^T c_t^T + P_t^N c_t^N$, subject to:

$U(c_t^T, c_t^N) = [\Omega (c_t^T)^{-\mu} + (1-\Omega)(c_t^N)^{-\mu}]^{-\frac{1}{\mu}}$ = 1, where UC is the budget of the household for obtaining one unit of utility; c_t^T and c_t^N are the shares of one unit composite utility. The first order conditions generate the following costs function: UC=$[\Omega^{\frac{1}{1+\mu}} (P_t^T)^{\frac{\mu}{1+\mu}} + (1-\Omega)^{\frac{1}{1+\mu}} (P_t^N)^{\frac{\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}$. In the perfect competitive equilibrium, the unit cost of obtaining the composite consumption goods equals the price of the goods. Hence, in the long-run, the price of the goods should take the form of: $P_t(P_t^T, P_t^N) = [\Omega^{\frac{1}{1+\mu}} (P_t^T)^{\frac{\mu}{1+\mu}} + (1-\Omega)^{\frac{1}{1+\mu}} (P_t^N)^{\frac{\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}$.

By substituting the price equation into the real exchange rate expression e_t with q_t denoting the nominal exchange rate, we can see:

$$e_t = q_t \frac{[\Omega^{*\frac{1}{1+\mu^*}} (P_t^T *)^{\frac{\mu^*}{1+\mu^*}} + (1-\Omega^*)^{\frac{1}{1+\mu^*}} (P_t^N *)^{\frac{\mu^*}{1+\mu^*}}]^{\frac{1+\mu^*}{\mu^*}}}{[\Omega^{\frac{1}{1+\mu}} (P_t^T)^{\frac{\mu}{1+\mu}} + (1-\Omega)^{\frac{1}{1+\mu}} (P_t^N)^{\frac{\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}}, \tag{13}$$

where * denotes the foreign country. Multiplying the numerator and denominator of equation (13) by $\frac{[(P_t^T *)^{\frac{-\mu^*}{1+\mu^*}}]^{\frac{1+\mu^*}{\mu^*}}}{[(P_t^T *)^{\frac{-\mu^*}{1+\mu^*}}]^{\frac{1+\mu^*}{\mu^*}}}$ and $\frac{[(P_t^T)^{\frac{-\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}}{[(P_t^T)^{\frac{-\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}}$, respectively, we have:

$$e_t = \frac{[\Omega^{*\frac{1}{1+\mu^*}} + (1-\Omega^*)^{\frac{1}{1+\mu^*}} (p_t^N *)^{\frac{\mu^*}{1+\mu^*}}]^{\frac{1+\mu^*}{\mu^*}}}{[\Omega^{\frac{1}{1+\mu}} + (1-\Omega)^{\frac{1}{1+\mu}} (p_t^N)^{\frac{\mu}{1+\mu}}]^{\frac{1+\mu}{\mu}}}, \tag{14}$$

where in the Balassa and Samuelson framework it is assumed that the PPP holds for tradable goods; we denote $\frac{P_t^N *}{P_t^T *}$ as $p_t^N *$ and $\frac{P_t^N}{P_t^T}$ as p_t^N . Eq. (14) demonstrates that home's real exchange rate against foreign depends only on the internal relative prices of nontradable goods.

2F. Regression Equations

Referring back to eq. (1), (11), (12), and (14), we can transform them into panel regression equations, respectively, as:

$$\ln p_{jt}^N = \delta_{0j} + \delta_1 \ln(\frac{K_{jt}^T}{Y_{jt}^T}) + \delta_2 \ln(\frac{K_{jt}^N}{Y_{jt}^N}) + \delta_3 \ln \theta_{jt}^T + \delta_4 \ln \theta_{jt}^N + \varepsilon_{jt} \tag{I}$$

$$\ln p_{jt}^N = \gamma_{0j} + \gamma_1 \ln(\frac{K_{jt}^T}{Y_{jt}^T}) + \gamma_2 \ln \theta_{jt}^T + \gamma_3 \ln \theta_{jt}^N + \varepsilon_{jt} \tag{II}$$

$$\ln p_{jt}^N = \eta_{0j} + \eta_1 \ln(\frac{I_{jt}^T}{Y_{jt}^T}) + \eta_2 \ln \theta_{jt}^T + \eta_3 \ln \theta_{jt}^N + \varepsilon_{jt} \tag{III}$$

$$\ln e_{jt} = \zeta_{oj} + \zeta_1 (\ln p_{jt}^{N*} - \ln p_{jt}^N) + \varepsilon_{jt} \tag{IV}$$

$(j=1,...,J, t=1,...,T)$.

For eq. (I), since labour productivity is a monotonic transformation of the capital-output ratio,[12] the relative price of nontradable goods is in line with the relative labour productivity of the tradable sector. Consequently, the coefficient on $\ln(\frac{K_{jt}^T}{Y_{jt}^T})$, which is δ_1, needs to be positive, and the coefficient on $\ln(\frac{K_{jt}^N}{Y_{jt}^N})$, which is δ_2, needs to be negative, if the Balassa-Samuelson effect holds.

Usually we restrict δ_3=-δ_4, and say that the coefficient on $(\ln\theta_{jt}^T - \ln\theta_{jt}^N)$ should be positive if the theory holds. Alternatively, we may estimate eq. (I) directly. The theory requires δ_3 and δ_4 to be positive and negative, respectively. By the same reasoning, we would expect that γ_2 in eq. (II), to be positive, and γ_3, to be negative. Also, η_2 and η_3, both in eq. (III), need to be positive and negative respectively.

Following Kravis, Heston, and Summers (1983) and Stockman and Tesar (1995),[13] we assume that tradables are relatively labour intensive, that is, $\alpha_T > \alpha_N$. Then, under such a condition, both γ_1 and η_1 should be negative.

The theory suggests that the real exchange rate and relative price differential between

---

[12] $\frac{Y_t^i}{N_t^i} = (A_t^i)^{\frac{1}{\alpha_i}} (\frac{K_t^i}{Y_t^i})^{\frac{1-\alpha_i}{\alpha_i}} X_t^i,\ i = T, N$.

[13] They find that labour's share of the income generated in the tradable sector is greater than that in the nontradable sector in their sample.

home and foreign countries are negatively correlated.[14] Hence, the coefficient on relative price differential, which is ζ_1 in eq. (IV), is expected to be negative. [15]

3. DATA AND VARIABLE CONSTRUCTION

Our main contribution to the literature is the construction of a large sectoral database for industrial analysis, which is primarily based on the China Statistical Yearbook, and includes eighteen fundamental variables, ten intermediate ones, and final eight variables across five sectors and thirty regions in China. Therefore the details of the data and the construction of the empirical counterparts to the theoretical variables should be put more in relief.

The nominal exchange rate is the annual average rate that is calculated based on monthly averages, in Chinese RMB yuan per U.S. dollar, from the IMF's *International Financial Statistics (IFS)*. Following such a definition, a decrease in the nominal exchange rate implies appreciation. The real exchange rate is defined as the bilateral real exchange rate between each region of China and the United States, adjusted to the difference in the GDP deflators of each region and the United States. The regional GDP deflator is the ratio of nominal to real GDP index (2000=1000)[16] for each region.

We follow De Gregorio and Wolf (1994) classification scheme, which classify sectors on the basis of export shares in output for the whole sample of regions with a cut-off

---

[14] A fall in home's real exchange rate against foreign implies appreciation due to the way we define the nominal exchange rate.

[15] Eq. (IV) has been empirically examined by, in particular, Chinn (2000) for Asia Pacific countries, and Vikas and Ogaki (1999) for several exchange rates.

[16] The real GDP index is obtained through the GDP index with the preceding year treated as 100.

point of 10% to delineate nontradables. The 10% threshold classifies the Chinese agriculture (farming, forestry, animal husbandry, and fishery) and industry (excavation, manufacturing, production and supply of power, gas, and water) as the tradable sector, and the remaining construction, transportation, storage, postal and telecommunications services, wholesale, retail trade, and catering services as the nontradable sector.

The sectoral prices (2000=100) for each region are the ratio of the nominal to real GDP index,[17] both at 2000 constant prices for each sector and region. The relative price differential is calculated as the difference in the relative prices of each Chinese region and the United States.

The TFPs are constructed from real GDP, capital stock, employment, and factor returns. The gross output value is the sum of the current value of final products produced in a given sector during a given period with the value of intermediate goods double counted (CSYB). Due to the lack of data on sectoral capital stock, all total capital is approximated through investment,[18] except the one for industry from 1993 to 2002, which is available and refers to "the capital received by the industrial enterprises from investors that could be used as operational capitals for a long period" (CSYB). Total employment, according to the definition given by the CSYB, is "the

[17] The sectoral GDP index (2000=100) is calculated through the real index of GDP (preceding year=100) in tradable and nontradable sectors, which is obtained using the fractions representing the composition of overall GDP and real GDP index by region and by individual sector.

[18] Investment is the capital construction investment in "new projects, including construction of a new facility, or an addition to an existing facility, and the related activities of the enterprises, institutions or administrative units mainly for the purpose of expanding production activity, covering only projects each with a total investment of 500,000 RMB yuan and over" (CSYB).

number of staff and workers, which refers to a literal translation of the Chinese term ‘zhigong’ that includes employees of state-owned units in urban and rural areas (including government agencies), of collective-owned units in urban areas, of other ownership units in urban areas, and of state-collective joint ownership.” Wages necessary to the construction of factor returns are the total wage bills of staffs and workers, which are also drawn from CSYB.

On the basis of the current OECD’s *STructural ANalysis (STAN)* industry list, the De Gregorio and Wolf (1994) 10% threshold classifies the U.S. agriculture, hunting, forestry and fishing, mining and quarrying, total manufacturing, electricity, gas, and water supply sectors as tradables, and the remaining construction, wholesale and retail trade, restaurants and hotels, transport, storage, and communication sectors as nontradables. The U.S. tradable and nontradable price deflators are constructed by dividing the nominal value added by the real value added (2000=100) for each sector, as reported in OECD’s *Annual National Accounts – Main Aggregates* under the code VALU and VALUK respectively.

4. EMPIRICAL RESULTS

In this section, we employ recently developed non-stationary panel data techniques that allow us to test the extended Asea and Mendoza (1994) model on Chinese Regional Data.

4A. Static panel data estimation

Tables 1 - 8 provide the estimates of eq. (I), (II), (III) and (IV) based on the pooled regression, OLS on differences, the least squares dummy variables (LSDV) regression using individual dummies in the OLS regression, the within estimates replacing y and

W by subtracting the means of each time series, the between estimates replacing y and W by the individual means, the feasible generalised least squares (GLS) estimates replacing y and W by deviations from weighted time means, the GLS using OLS residuals, and the maximum likelihood estimates (MLE) obtained by iterating the GLS procedure (Baltagi 1995).[19]

Among these static panel models, the Balassa-Samuelson proposition is supported by the data in the Total, LSDV, within-groups, GLS using within/between groups, GLS using OLS residuals, and MLE models of eq. (I) and (IV) – all coefficients are statistically significant and of the expected signs. When looking at the outputs for eq. (I), the magnitudes of the sectoral TFP coefficients suggest that the data is able to support our extended model in that the homogeneity restrictions on TFPs across sectors, are rejected. One thing that we should be aware of is that the residuals in those regressions do not pass the diagnostic tests so well – they all fail the AR(1) test, however, they do not pass the AR(2) tests.

When estimating eq. (II), in six out of the eight cases, the coefficients of the sectoral TFPs, that is, γ_2 and γ_3, are significant and of the correct signs. However, the coefficient of the tradable capital-output ratio, γ_1, remains positive in almost all cases,

---

[19] The linear model is given by: $y_{jt} = x_{jt}'\gamma + \lambda_t + \eta_j + v_{jt} \quad (j = 1,...,J, t = 1,...,T)$, where λ_t is the time effect, η_j is the fixed individual effect, x_{jt} is a $k \times 1$ vector of time-varying explanatory variables assumed to be strictly exogenous, v_{jt} is a vector of the independently and identically distributed errors. Stacking the data for an individual according to time, and then stacking all individuals, and combining the data into W= [X:D] yields $y = W\beta + v$.

which is inconsistent with the theory.

The results from estimating eq. (III) are less satisfactory. Although the coefficient of the nontradable TFP, η_3, appears significant and negative in almost all cases, the remaining coefficients are of the wrong sign, or not statistically different from zero. Residuals from most regressions do not pass the AR(2) tests so well.

4B. Dynamic panel data estimation

In this section, we estimate eq. (I) - (IV) in levels, using one- and two-step GMM (Arellano and Bond 1991) and combined GMM estimation (Arellano and Bover 1995; Blundell and Bond 1998). The standard errors and tests are based on the robust variance matrix. In order to determine the proper lag length, we estimate equations with different combinations of the lag structure of the x_{jt} matrix. Among our experiments, we choose to look at the results where residuals pass both the Sargan[20] and AR(2) test, and fail the AR(1) test.[21]

When we estimate eq. (I), we choose the results generated by the two-step GMM estimation with one lag on relative price, tradable capital-output ratio, and

---

[20] The Sargan (1958, 1988) test tests the over-identifying restrictions. That is, if A_J is optimal for any given Z_j, then under the null hypothesis that the instruments in Z are exogenous (i.e. uncorrelated with the individual effect η_1), the test statistic is $(\sum_{j=1}^{J} \hat{v}_j *' z_j) A_J (\sum_{j=1}^{J} z_j ' \hat{v}_j *) \sim \chi_r^2$.

[21] If the AR(1) model is mean-stationary, then Δy_{jt} are uncorrelated with η_1, which suggests that $\Delta y_{j,t-1}$ can be used as instruments in the levels equations (Arellano and Bover 1995; Blundell and Bond 1998).

nontradable TFP, and two lags on nontradable capital-output ratio and tradable TFP, respectively (Table 9). Under such specifications of instruments in GMM estimators, the residuals pass all diagnostic tests well. All the estimated coefficients have the expected signs; although the tradable capital-output ratio (δ_1) and TFP (δ_3) are insignificant.

We follow the same lag selection procedure described above to estimate eq. (II) - (IV) (Table 9). The coefficient that appears consistent with the theory is the one on nontradable TFP, which is negative and statistically different from zero throughout eq. (II) and (III). The remaining coefficients, such as the ones on investment-output ratio (η_1) and on relative price differential (ζ_1), do not have the expected signs; the coefficients of tradable TFP in eq. (II) and (III), are either insignificant (p-value=0.23), or negative (-0.04).

For combined GMM estimation (Table 10), the results are generally similar to the GMM estimation with one exception. The estimated coefficient on relative price differential, which is ζ_1 in eq. (IV), is negative and significant with residuals passing all the diagnostic tests. Thus, China is observed a correct direction in the change in the relative prices - assuming that the law of one price for tradable goods hold, then, for China, during its fast growing 1985 – 2000 period, the higher the relative price of nontradable goods, the lower the real exchange rate would become.

5. CONCLUSION

The paper extends the Asea and Mendoza (1994) setting to consider asymmetric productivity shocks across sectors. Testing this model on Chinese Regional Data aid of non-stationary panel data techniques, it shows that eq. (I) of our extended Asea and

Mendoza (1994) model is a reasonable empirical representation of the Chinese Balassa-Samuelson effect. The static panel data model seems mis-specified as it left out all the dynamics, and the best results are, as expected, from the dynamic one. This can be seen by looking at the Sargan test, which suggests that the instruments are exogenous. In fact, Hall and Urga (1998) show that when T is small and J is large, the GMM estimator is an efficient estimator, especially when taking the first differences or orthogonal deviations to eliminate the fixed effects. In addition, the combined two-step GMM estimation shows that China has managed to keep its real exchange rate appreciated while its growth rate is respectably high.

REFERENCES

Arellano, M. and Bond, S. R., 1991, "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies*, 58, 277-297.

Arellano, M. and Bover, O., 1995, "Another Look at the Instrumental Variables Estimation of Error-Components Models," *Journal of Econometrics*, 68, 29-51.

Asea, P. and Mendoza, E., 1994, "The Balassa-Samuelson Model: a General Equilibrium Appraisal," *Review of International Economics,* 2, 244-267.

Balassa, B., 1964, "The Purchasing Power Parity Doctrine: A Reappraisal," *Journal of Political Economy,* 72, 584-596.

Baltagi, B. H., 1995, *Econometric Analysis of Panel data*, John Wiley and Sons.

Bergstrand, J. H., 1991, "Structural Determinants of Real Exchange Rates and National Price Levels: Some Empirical Evidence," *American Economic Review*, 81(1), 325-34.

Blundell, R. W. and Bond, S. R., 1998, “Initial Conditions and Moment Restrictions in Dynamic Panel Data Models,” *Journal of Econometrics*, 87, 115-143.

Canzoneri, M. B., Cumby, R. E., and Diba, B., 1999, “Relative Labour Productivity and the Real Exchange Rate in the Long-Run: Evidence for a Panel of OECD Countries,” *Journal of International Economics*, 47, 245-266.

China Statistical Yearbook, China Statistical Bureau, various issues.

Chinn, M., 1995, “Paper Pushers or Paper Money? Empirical Assessment of Fiscal and Monetary Models of Real Exchange Rates,” *Journal of Policy Modeling*, 19, 51-78.

Chinn, M., 1996, “Whither the Yen? Implications of an Intertemporal Model of the Yen/Dollar Rate,” *Mimeo*

Chinn, M., 2000, “The Usual Suspects? Productivity and Demand Shocks and Asia-Pacific Real Exchange Rates,” *Review of International Economics*, 8(1), 20-43.

Chinn, M. and Johnston, L., 1996, "Real Exchange Rate Levels, Productivity and Demand Shocks: Evidence from a Panel of 14 Countries," *IMF Working Papers,* 97/66.

De Gregorio, J., Giovannini, A., and Krueger, T., 1994, "The Behavior of Nontradable-Goods Prices in Europe: Evidence and Interpretation," *Review of International Economics*, 2(3), 284-305.

De Gregorio, J., Giovannini, A., and Wolf, H., 1994, "International Evidence on Tradables and Non-tradables Inflation," *European Economic Review,* 38, 1255-1244.

De Gregorio, J. and Wolf, H., 1994, "Terms of Trade, Productivity, and the Real Exchange Rate," *NBER Working Paper,* 4807.

Drine, I. and Rault, C., 2002, "Does the Balassa-Samuelson Hypothesis Hold for Asian Countries? An Empirical Analysis using Panel Data Cointegration Tests," *William Davidson Institute Working Paper Series*, 504, University of Michigan Stephen M. Ross Business School (September 2002).

Drine, I. and Rault, C., 2003, "Do Panel Data Permit the Rescue of the Balassa-Samuelson Hypothesis for Latin American Countries?" *Applied Economics*, 35, 2003, 351-59.

Evans, C., 1992, "Productivity Shocks and Real Business Cycles," *Journal of Monetary Economics*, 29, 191-208.

Froot, K. and Rogoff, K., 1991, "The EMS, the EMU and the Transition to a Common Currency," *NBER Macroeconomics Annual* 1991 (MIT Press), 269-317.

Guillaumont Jeanneney, S. and Hua, P., 2002, “The Balassa–Samuelson Effect and Inflation in the Chinese Provinces,” *China Economic Review,* 13 (2002), 134–160.

Hall, S. and Urga, G., 1998, “Stochastic Common Trends and Long-run Relationships in Heterogeneous Panels,” *Discussion paper*, Centre for Economic Forecasting, London Business School, 27-95.

Hsieh, D., 1982, “The Determination of the Real Exchange Rate: The Productivity Approach,” *Journal of International Economics,* 12 (2) (May), 355-362.

Ito, T., Isard, P., and S. Symansky, 1997, “Economic Growth and Real Exchange Rate: An Overview of the Balassa-Samuelson Hypothesis in Asia,” *NBER working paper* 5979.

Kravis, I. B., Heston A. W., and Summers R., 1983, “The Share of Services in Economic Growth,” in F. G. Adams and B. Hickman (eds.), *Essays in Honour of Lawrence R. Klein*, MIT Press, Cambridge, MA, 1983.

Marston, R., 1987, “Real Exchange Rates and Productivity Growth in the United States in Japan,” in Arndt, Sven and J. David Richardson (eds), *Real-Financial Linkages among Open Economies*, MIT Press, 1987

Marston, R., 1990, “Systematic Movements in Real Exchange Rates in the G-5: Evidence on the Integration of Internal and External Markets,” *Journal of Banking and Finance,* 14(5) (November), 1023-1044.

Micossi, S. and Milesi-Ferretti, G. M., 1994, "Real Exchange Rates and the Prices of Nontradable Goods," *IMF Working Paper,* 94/19 (February).

Obstfeld, M., 1993, "Model Trending Real Exchange Rates," *Working Paper,* C93-011, Center for International and Development Economics Research, University of California, Berkeley.

Rogoff, K., 1992, "Traded Goods Consumption Smoothing and the Random Walk Behavior of the Real Exchange Rate," *Bank of Japan Monetary and Economic Studies,* 10(2), 1-29. Also, *NBER Working Paper,* 4119.

Samuelson, P., 1964, "Theoretical Notes on Trade Problems," *Review of Economics and Statistics,* 46, 145-154.

Sargan, J. D., 1958, "The Estimation of Economic Relationships Using Instrumental Variables," *Econometrica*, 26, 393-415.

Sargan, J. D., 1988, "Testing for Misspecification after Estimating Using Instrumental Variables," in E. Maasoumi, ed., *Contributions to Econometrics: John Denis Sargan,* vol.1. Cambridge, Cambridge University Press.

Solow, R. M., 1956, "A Contribution to the Theory of Economic Growth," *Quarterly Journal of Economics,* 70 (February), 65-94. Reprinted in Stiglitz and Uzawa (1969).

Stockman, A. C. and Tesar L. L., 1995, "Tastes and Technology in a Two-Country Model of the Business Cycle: Explaining International Comovements," *American*

Economic Review, American Economic Association, 85(1), pages 168-85, March.

Strauss, J., 1995, "Real Exchange Rates, PPP and the Relative Price of Nontraded Goods," *Southern Economic Journal*, 61(4) (April), 991-1005.

Vikas, K. and Ogaki, M., 1999, "Real Exchange Rates and Nontradables: A Relative Price Approach," *Journal of Empirical Finance,* 6 (April 1999), 193-215.

TABLES

Table 1 Pooled (Total) regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	0.75** (0.01)	γ_1	0.15 (0.63)	η_1	-0.15** (0.00)	ς_1	-0.58** (0.00)
δ_2	-0.78** (0.00)	γ_2	-0.04 (0.91)	η_2	-0.31** (0.00)	- -	- -
δ_3	0.60* (0.08)	γ_3	-0.07 (0.16)	η_3	-0.10** (0.03)	- -	- -
δ_4	-0.92** (0.00)	- -	- -	- -	- -	- -	- -
Constant	0.79** (0.00)	Constant	0.37** (0.05)	Constant	0.13 (0.41)	Constant	2.54** (0.00)
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	4.04** (0.00)	AR(1) test	4.44** (0.00)	AR(1) test	4.29** (0.00)	AR(1) test	25.13** (0.00)
AR(2) test	3.97** (0.00)	AR(2) test	4.14** (0.00)	AR(2) test	4.03** (0.00)	AR(2) test	16.01** (0.00)

Table 2 OLS on differences regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	0.44 (0.32)	γ_1	0.70** (0.05)	η_1	0.19** (0.01)	ζ_1	-0.02 (0.69)
δ_2	-0.88** (0.00)	γ_2	0.66* (0.08)	η_2	-0.04** (0.00)	- -	- -
δ_3	0.42 (0.36)	γ_3	-0.31** (0.00)	η_3	-0.28** (0.00)	- -	- -
δ_4	-1.06** (0.00)	- -	- -	- -	- -	- -	- -
Constant	-0.02** (0.01)	Constant	-0.00 (0.69)	Constant	-0.01** (0.01)	Constant	-0.07** (0.00)
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	-1.89* (0.06)	AR(1) test	-1.61 (0.11)	AR(1) test	-1.23 (0.22)	AR(1) test	-3.83** (0.00)
AR(2) test	1.96 (0.05)	AR(2) test	0.67 (0.51)	AR(2) test	0.61 (0.54)	AR(2) test	-2.73** (0.01)

Notes: 1) The figures in parentheses refer to p-values; 2) Statistical significance at 5% and 10% levels are denoted by ** and * respectively.

Table 3 Least squares dummy variables (LSDV) regression using individual dummies in the OLS regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	1.19** (0.00)	γ_1	0.95** (0.00)	η_1	0.09 (0.16)	ζ_1	-0.68** (0.00)
δ_2	-0.68** (0.00)	γ_2	0.75** (0.01)	η_2	-0.36** (0.00)	- -	- -
δ_3	1.05** (0.02)	γ_3	-0.20** (0.00)	η_3	-0.13** (0.01)	- -	- -
δ_4	-0.87** (0.00)	- -	- -	- -	- -	- -	- -
Constant	0.54** (0.00)	Constant	0.07 (0.67)	Constant	0.86** (0.00)	Constant	2.64** (0.00)
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	3.70** (0.00)	AR(1) test	4.19** (0.00)	AR(1) test	4.27** (0.00)	AR(1) test	19.47** (0.00)
AR(2) test	3.21** (0.00)	AR(2) test	3.76** (0.00)	AR(2) test	3.15** (0.00)	AR(2) test	10.05** (0.00)

Table 4 Within-groups regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	1.19** (0.00)	γ_1	0.95** (0.00)	η_1	0.09 (0.16)	ζ_1	-0.68** (0.00)
δ_2	-0.68** (0.00)	γ_2	0.75** (0.01)	η_2	-0.36** (0.00)	- -	- -
δ_3	1.05** (0.02)	γ_3	-0.20** (0.00)	η_3	-0.13** (0.01)	- -	- -
δ_4	-0.87** (0.00)	- -	- -	- -	- -	- -	- -
Constant	- -	Constant	- -	Constant	- -	Constant	- -
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	3.70** (0.00)	AR(1) test	4.19** (0.00)	AR(1) test	4.27** (0.00)	AR(1) test	20.05** (0.00)
AR(2) test	3.21** (0.00)	AR(2) test	3.76** (0.00)	AR(2) test	3.15** (0.00)	AR(2) test	10.82** (0.00)

Notes: 1) The figures in parentheses refer to p-values; 2) Statistical significance at 5% and 10% levels are denoted by ** and * respectively.

Table 5 Between-groups regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	0.25 (0.68)	γ_1	-0.00 (1.00)	η_1	0.09 (0.66)	ζ_1	-0.08 (0.52)
δ_2	-0.40 (0.22)	γ_2	-0.02 (0.98)	η_2	0.12 (0.71)	- -	- -
δ_3	0.19 (0.80)	γ_3	-0.03 (0.78)	η_3	-0.04 (0.72)	- -	- -
δ_4	-0.48 (0.21)	- -	- -	- -	- -	- -	- -
Constant	0.47 (0.32)	Constant	0.11 (0.77)	Constant	0.23 (0.45)	Constant	2.66** (0.00)
Trend	No	Trend	no	Trend	no	Trend	no
AR(1) test	- -	AR(1) test	- -	AR(1) test	- -	AR(1) test	- -
AR(2) test	- -	AR(2) test	- -	AR(2) test	- -	AR(2) test	- -

Table 6 GLS using within/between-groups regression

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	1.14** (0.00)	γ_1	0.80** (0.00)	η_1	-0.01 (0.83)	ζ_1	-0.63** (0.00)
δ_2	-0.70** (0.00)	γ_2	0.59** (0.00)	η_2	-0.34** (0.00)	- -	- -
δ_3	1.00** (0.00)	γ_3	-0.18** (0.00)	η_3	-0.13** (0.00)	- -	- -
δ_4	-0.88** (0.00)	- -	- -	- -	- -	- -	- -
Constant	0.65** (0.00)	Constant	0.31** (0.00)	Constant	0.69** (0.00)	Constant	2.53** (0.00)
Trend	No	Trend	no	Trend	no	Trend	no
AR(1) test	13.96** (0.00)	AR(1) test	17.94** (0.00)	AR(1) test	15.64** (0.00)	AR(1) test	22.84** (0.00)
AR(2) test	9.22** (0.00)	AR(2) test	10.64** (0.00)	AR(2) test	7.33** (0.00)	AR(2) test	13.51** (0.00)

Notes: 1) The figures in parentheses refer to p-values; 2) Statistical significance at 5% and 10% levels are denoted by ** and * respectively.

Table 7 GLS using OLS residuals

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	1.13** (0.00)	γ_1	0.79** (0.00)	η_1	-0.01 (0.65)	ζ_1	-0.64** (0.00)
δ_2	-0.70** (0.00)	γ_2	0.59** (0.00)	η_2	-0.34** (0.00)	- -	- -
δ_3	0.99** (0.00)	γ_3	-0.18** (0.00)	η_3	-0.13** (0.00)	- -	- -
δ_4	-0.88** (0.00)	- -	- -	- -	- -	- -	- -
Constant	0.65** (0.00)	Constant	0.31** (0.00)	Constant	0.66** (0.00)	Constant	2.52** (0.00)
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	14.22** (0.00)	AR(1) test	18.00** (0.00)	AR(1) test	15.84** (0.00)	AR(1) test	22.03** (0.00)
AR(2) test	9.49** (0.00)	AR(2) test	10.70** (0.00)	AR(2) test	7.58** (0.00)	AR(2) test	12.60** (0.00)

Table 8 Maximum likelihood estimates (MLE) obtained by iterating the GLS procedure

Coefficient	Eq. (I)	Coefficient	Eq. (II)	Coefficient	Eq. (III)	Coefficient	Eq. (IV)
δ_1	1.15** (0.00)	γ_1	0.88** (0.00)	η_1	0.04 (0.26)	ζ_1	-0.65** (0.00)
δ_2	-0.69** (0.00)	γ_2	0.68** (0.00)	η_2	-0.35** (0.00)	- -	- -
δ_3	1.01** (0.00)	γ_3	-0.19** (0.00)	η_3	-0.13** (0.00)	- -	- -
δ_4	-0.88** (0.00)	- -	- -	- -	- -	- -	- -
Constant	0.64** (0.00)	Constant	0.30** (0.00)	Constant	0.84** (0.00)	Constant	2.52** (0.00)
Trend	no	Trend	no	Trend	no	Trend	no
AR(1) test	13.44** (0.00)	AR(1) test	16.15** (0.00)	AR(1) test	14.51** (0.00)	AR(1) test	21.89** (0.00)
AR(2) test	8.68** (0.00)	AR(2) test	8.66** (0.00)	AR(2) test	5.99** (0.00)	AR(2) test	12.44** (0.00)

Notes: 1) The figures in parentheses refer to p-values; 2) Statistical significance at 5% and 10% levels are denoted by ** and * respectively.

Table 9 One- or two-step GMM regression

Coefficient	Eq. (I)	Lag	Coefficient	Eq. (II)	Lag	Coefficient	Eq. (III)	Lag	Coefficient	Eq. (IV)	Lag
δ_1	0.56 (0.40)	1	γ_1	0.60 (0.18)	1	η_1	0.21** (0.00)	1	ζ_1	0.08* (0.08)	1
δ_2	-0.95** (0.00)	2	γ_2	0.57 (0.23)	2	η_2	-0.04** (0.01)	1	- -	- -	- -
δ_3	0.56 (0.42)	2	γ_3	-0.31** (0.00)	1	η_3	-0.19** (0.07)	2	- -	- -	- -
δ_4	-0.98** (0.00)	1	-	-	-	-	-	-	-	-	-
Constant	0.03 (0.75)		Constant	-0.00 (0.99)		Constant	0.01 (0.74)		Constant	-0.05** (0.00)	
Trend	no		Trend	no		Trend	yes		Trend	no	
1 or 2-step	2-step		1 or 2-step	2-step		1 or 2-step	1-step		1 or 2-step	2-step	
Sargan test	25.19 (1.00)		Sargan test	28.55 (1.00)		Sargan test	104.00 (0.48)		Sargan test	29.95 (1.00)	
AR(1) test	-1.62* (0.10)		AR(1) test	-1.23 (0.22)		AR(1) test	-2.25** (0.02)		AR(1) test	-3.64** (0.00)	
AR(2) test	1.41 (0.16)		AR(2) test	-0.02 (0.98)		AR(2) test	-1.43 (0.15)		AR(2) test	-0.29 (0.77)	

Table 10 One- or two-step combined GMM regression

Coefficient	Eq. (I)	Lag	Coefficient	Eq. (II)	Lag	Coefficient	Eq. (III)	Lag	Coefficient	Eq. (IV)	Lag
δ_1	0.15 (0.56)	1	γ_1	0.49 (0.13)	1	η_1	0.20** (0.00)	1	ζ_1	-0.12** (0.04)	1
δ_2	-0.64** (0.00)	1	γ_2	0.47 (0.19)	1	η_2	-0.04** (0.04)	1	- -	- -	- -
δ_3	0.08 (0.74)	1	γ_3	-0.27** (0.00)	1	η_3	-0.12* (0.07)	1	- -	- -	- -
δ_4	-0.71** (0.00)	1	-	-	-	-	-	-	-	-	-
Constant	0.22** (0.02)		Constant	0.01 (0.86)		Constant	0.15** (0.00)		Constant	0.28** (0.00)	
Trend	yes		Trend	no		Trend	yes		Trend	no	
1 or 2-step	1-step		1 or 2-step	1-step		1 or 2-step	1-step		1 or 2-step	2-step	
Sargan test	337.00 (1.00)		Sargan test	495.30 (0.19)		Sargan test	272.60 (1.00)		Sargan test	29.94 (1.00)	
AR(1) test	-1.63* (0.10)		AR(1) test	-1.23 (0.21)		AR(1) test	-1.79* (0.07)		AR(1) test	-3.83** (0.00)	
AR(2) test	0.82 (0.41)		AR(2) test	1.03 (0.30)		AR(2) test	0.59 (0.56)		AR(2) test	-0.07 (0.94)	

Notes: 1) The figures in parentheses refer to p-values; 2) Statistical significance at 5% and 10% levels are denoted by ** and * respectively.