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REAL EXCHANGE RATES IN THE LONG AND SHORT RUN: A PANEL CO-INTEGRATION APPROACH

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Resumen

El presente artículo pretende enfrentar los problemas de la literatura empírica del tipo de cambio real de largo plazo (pobre cobertura temporal de los datos y pobre desempeño de raíces unitarias en series de tiempo) mediante el uso de técnicas de cointegración para datos de panel. Usando datos anuales para 67 países durante el periodo 1996-97, hallamos evidencia de cointegración entre el tipo de cambio y sus fundamentos (activos externos netos, productividad y términos de intercambio) para la muestra completa y para sub-muestras de acuerdo a niveles de ingreso y controles de capital. El análisis de robustez de nuestros parámetros de largo plazo revela: (i) la estimación de una relación de largo plazo común para todos los países entre el tipo de cambio y sus fundamentos es válida sólo para países con niveles de ingreso alto y para países con bajos controles de capital, (ii) existe un corte estructural en la relación de largo plazo en 1973 para estas mismas sub-muestras. Finalmente, desviaciones del tipo de cambio real de equilibrio son persistentes con una vida media estimada entre 2.8 y 5 años, consistente con el intervalo de [2.5-5.0] años hallado en la literatura (Murray and Papell, 2002).

Abstract

The empirical literature on long-run real exchange rate behavior has shown mixed evidence due to problems involving the lack of long time series data and the low power of time-series unit root tests in small samples. The main objective of the present paper is to tackle these empirical issues by applying the recently developed panel cointegration techniques to the long-run real exchange rate equation implied by our model. Using annual data for 67 countries over the 1966-97, we find that the cointegrating relationship between the real exchange rate, the ratio of net foreign assets to GDP, the relative Home to Foreign productivity of the traded and non-traded sector, and the terms of trade is valid in the long run. This result holds for all sub-sample of countries (whether they are classified by income per capita or capital controls). Furthermore, our coefficient estimates are consistent with the theoretical values implied by the calibrated parameters of preferences and technology in Stockman and Tesar (1995). Robustness checks reveal that: (i) "pooling" the data to obtain a common long-run equilibrium relationship across countries is valid for the samples of countries with high income and low capital controls, (ii) the oil shock crisis in 1973 represents a structural change for these sub-samples. Finally, deviations from the equilibrium are large and persistent with half-life estimates (between 2.8 and 5) consistent with the consensus interval of 2.5-5 found in the literature (Murray and Papell, 2002).

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1. Introduction

The Purchasing Power Parity (PPP), as a benchmark for evaluating long real exchange rate behavior, has been revitalized by recent improvements in methods of empirical testing and the availability of data sets with longer time series coverage (Froot and Rogoff, 1995). However, persistent deviations from *PPP* have been empirically observed (Wei and Parsley, 1995) and theoretically linked to the evolution of fundamentals in the economy (Lucas, 1982; Stockman, 1987), with real exchange rate fluctuations being driven by real shocks that represent shifts in the relative prices consistent with the international equilibrium.

The empirical literature on long-run real exchange rate behavior has shown mixed evidence. This could be mainly attributed to the lack of long time series data and the low power of time-series unit root tests in small samples to distinguish between non-stationary and stationary but highly persistent processes (Canzoneri *et al.*, 1999). To tackle the empirical problems mentioned above, we use the recently developed panel cointegration techniques (Kao, 1999; Kao and Chiang, 1999; Pedroni, 1999, Phillips and Moon, 1999) to estimate the long run equilibrium relationship for the real exchange rate. Panel cointegration techniques will allow us to deal with non-stationary data for a heterogeneous panel of 67 countries with annual data over the period 1966-97. Finally, to interpret our results, we derive a long run real exchange rate equation from a simple *"real"* version of the Obstfeld-Rogoff (1995) model, which incorporates traded and non-traded goods.

As predicted in our model, we find evidence of a long-run cointegration relationship between the real exchange rate, and the ratio of net foreign assets to GDP, relative (home to foreign) productivity of the traded and non-traded sectors, and the terms of trade. This result holds for the full sample and all sub-samples of countries (classified by income per capita and by capital controls). After testing the robustness of our long run estimates, we find that: (1) *"pooling"* across individuals (cross-country parameter stability) is a valid assumption only for the samples of high-income countries and countries with low capital controls. (2) There is evidence of parameter stability over time, with the first oil shock crisis (1973) being a source of structural break for the full sample of countries as well as for the sample of high-income countries and countries with low capital controls. (3) After calibrating the theoretical long-run parameters in our real exchange equation with values for preferences and technology parameters (taken from Stockman and Tesar, 1995) we find that our estimated coefficients are in line with these theoretical values. However, the discrepancies between them are larger for the sample of high and upper-middle income countries, and depend on our assumption for the elasticity of intertemporal substitution.

Finally, we explore the half-life of real exchange rate deviations from the equilibrium. For the full sample of countries, we find that these deviations have an average decay rate of *21.2* percent per year, with an implied half-life of *3.3* years. On the other hand, the average half-life of these deviations for high-income countries is *2.87* years, which lies in the lower bound of the consensus interval of 2.5-5 year half-lives for PPP deviations found in the literature (Murray and Papell, 2002).

Although the literature has devoted some attention to the issue of equilibrium real exchange rates (Edwards, 1989; Faruqee, 1995; Balvers and Bergstrand, 1997; McDonald and Stein, 1999), we consider that our work complements and improves the existing literature in several aspects. First, we overcome the low power of the time-series unit roots and cointegration testing procedure by applying the recently developed panel unit roots and cointegration techniques. Second, we formulate a simple model of real exchange rate determination in the spirit of the new open economy macroeconomics that could be used as a benchmark for evaluating real exchange rate behavior in the long run. Third, we thoroughly test the parameter stability of our long run coefficient estimates across countries and over time (i.e. tests of country heterogeneity and structural change, respectively). Finally, we explore the short-run dynamics of deviations from the equilibrium real exchange rate and try to characterize the half-life of these deviations.

The rest of the paper is organized as follows. Section 2 presents the theoretical model that poses the long-run equilibrium real exchange rate equation, which also represents the fundamental equation of our empirical assessment. Section 3 discusses the data used in our empirical evaluation. Section 4 presents the panel data estimates for the long run real exchange rate. Section 5 introduces different robustness checks on our long-run estimates between and within countries as well as over time. Section 6 discusses the behavior of real exchange rates in the short-run, thus computing half-lives of equilibrium real exchange rate deviations. Finally, section 7 concludes.

2. The Theoretical Model

The present section develops a simple model of real exchange rate behavior. The basis of this model is the extended version of the Obstfeld and Rogoff (1995) model that incorporates traded and non-traded goods. In order to evaluate the fundamentals of the real exchange rate, we further assume that there is no money in the economy. In what follows, we state the basic assumption and implications of the model.[1]

2.1. A Basic Setup

Consider a two-country model with the non-traded sector being the locus of the monopoly and sticky price problems, and where the traded sector has a single homogeneous output that is priced in competitive world markets. Each representative agent of the Home country is endowed with a constant quantity of the traded good each period, $\overline{y}_T$, and has a monopoly power over one of the non-tradables goods $z \in [0,1]$. All producers reside in two countries, Home and Foreign. Home country consists of producers on the interval [0,n], whereas Foreign producers are located on (n,1]. We assume that all agents have similar preferences throughout the world over a real

[1]Lane and Milesi-Ferreti (2000) have formulated a similar approach to reassess the empirical evidence on the transfer effect for a sample of 64 countries (mostly, industrial and high and upper-middle income) over the 1970-96 period.

consumption index and work effort. Given the symmetry in preferences and budget constraints across agents, we solve the optimization problem for the representative national consumer-producer.

The intertemporal utility function of the typical Home agent *j* is given by:

$$U_t^j = \sum_{s=t}^{\infty} \beta^{s-t} \left[\frac{\sigma}{\sigma - 1} C_s^{1-\frac{1}{\sigma}} - \frac{\kappa}{2} y_{N,s}^2 \right] \tag{1}$$

where $\beta \in (0,1)$, and $\sigma, \kappa > 0$.[2] The consumption index, C, is an aggregate index of tradable and non-tradable consumption (C_N and C_T, respectively):

$$C_t = \left[\gamma^{1/\theta} C_{T,t}^{\frac{\theta-1}{\theta}} + (1-\gamma)^{1/\theta} C_{N,t}^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}} \tag{2}$$

with θ representing the elasticity of intra-temporal substitution (i.e. elasticity of substitution between traded and non-traded consumption). Agent *j* can invest in an internationally traded asset (denominated in units of the import good), and her flow budget constraint is given by:

$$F^j_{t+1} = (1+r_t)F^j_t + p_{Nt}(j)y_{Nt}(j) - P^X_{Tt}\bar{y}_{Tt} - P_t C^j_t \tag{3}$$

where F_t denotes real bonds (in units of the tradable good) that pay off a real return *r*, $p_{Nt}(j)$ is the price of the non-traded good produced by agent *j*, and P^X_{Tt} is the world price of the non-traded good. The consumer price index (CPI) for the Home country is given by:

$$P_t = \left[\gamma P_{T,t}^{1-\theta} + (1-\gamma) P_{N,t}^{1-\theta} \right]^{\frac{1}{1-\theta}} \tag{4}$$

with P_{Tt} and P_{Nt} being the prices of traded and non-traded goods at time *t*, respectively. In addition, the real exchange rate (Q_t) is defined as the ratio of domestic to foreign price consumer index,[3]

$$Q_t = \frac{P_t}{P_t^*} \tag{5}$$

Finally, the producer of non-traded goods face the following demand curve:

$$y^d_{N,t} = \left[\frac{p_{N,t}(j)}{P_{N,t}} \right]^{-\theta} C_N^A \tag{6}$$

where C_N^A represents Home's aggregate consumption of non-traded goods.

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[2]The final term in the period utility function, $-(\kappa/2)y_{N,s}^2$, captures the dis-utility experienced in producing more output. If the dis-utility from effort ℓ_N is given by $-\psi\, \ell_N$ and the production function is $y_N = A\ell_N^{\alpha}$ ($\alpha<1$), then $\kappa = 2\psi / A^{1/\alpha}$. If $\alpha = 0.5$, we have the output term in equation (1). Note that a rise in productivity A is captured in this model by a fall in κ.

[3]Note that the CPI-based real exchange rate is independent of the terms of trade in this model. Here, real exchange rates might be influenced by the terms of trade indirectly through wealth effects on the relative price of non-tradables.

To solve the agent's optimization problem, we maximize equation (1) subject to equations (3) and (6). The solution for the consumption and work effort paths might meet the following first-order conditions:

$$\frac{C_{T,t+1}}{C_{T,t}} = [b(1+r_{t+1})]^{s} \left(\frac{P_{T,t+1}}{P_{T,t}} \right)^{-q} \left(\frac{P_t}{P_{t+1}} \right)^{s-q} \tag{7}$$

$$\frac{C_{N,t}}{C_{T,t}} = \left(\frac{g}{1-g} \right) \left(\frac{P_{N,t}}{P_{T,t}} \right)^{-q} \tag{8}$$

$$y_{N,t}^{\frac{q+1}{q}} = \left(\frac{q-1}{qk} \right) C_t^{-1/s} \left(C_{N,t}^{A} \right)^{1/q} \left(\frac{P_{N,t}}{P_t} \right) \tag{9}$$

First, we find that consumption of tradables depends on the sequence of relative prices (i.e. the consumption-based real interest rate effect). Thus, if the aggregate price level relative to the price of tradables is currently low relative to its future value, then present consumption is preferred over future consumption as the consumption-based real interest rate is lower. However, it also encourages substitution from traded to non-traded goods. The former effect dominates if the intertemporal elasticity of substitution is greater than the intra-temporal elasticity of substitution (i.e. $\sigma > \theta$).

Second, the relationship between consumption of non-traded and traded goods is specified in equation (8), with θ being the elasticity of substitution between traded and non-traded goods. If the relative price is equal to one, the relative consumption of non-traded goods is larger, the smaller is the parameter γ.

Third, equation (9) postulates the equilibrium supply of non-traded goods. Note that the higher is the consumption index *C*, the lower is the level of production, as agents increase leisure in line with consumption of other goods.[4]

2.2. Approximate Solution

Consider the benchmark steady state in which all variables are constant.[5] We normalize the endowment of the traded good so that the relative price of non-traded goods in terms of traded goods P_N is equal to one. In addition, we assume that the price of the traded goods $P_{T,t}^{X}$ is equal to one. In this symmetric equilibrium, the steady state production and consumption of non-traded and traded goods are given by:[6]

[4]Note that the first-order conditions, eq.(7)-(9), and the period budget constraint, eq.(3), do not fully characterize the equilibrium. We also require the transversality condition: $\lim_{T \to \infty} \prod_{v=s}^{t+T} \left(\frac{1}{1+r_v} \right) F_{t+T+1}$.

[5]We assume that the stock of net foreign assets is zero, *i.e.* $nF_{t+1} + (1-n)F_{t+1}^{*} = 0$.

[6]Equation (10) finds that the less taxing is work effort (the smaller is κ), the larger the production of non-tradables will be in the steady state. In addition, equation (11) states that the larger is the weight placed on consumption of tradables in the utility function (i.e. the larger is γ), the larger is the ratio of traded to non-traded output.

$$Y_N = C_N = \left(\frac{q-1}{qk}\right)^{\frac{s}{1+s}} (1-g)^{\frac{1}{1+s}} \tag{10}$$

$$Y_T = C_T = \left(\frac{g}{1-g}\right) Y_N \tag{11}$$

Next, we take a log-linear approximation around the benchmark steady state. Let $\tilde{X} \equiv dX/X_0$ denote the percentage change relative to the benchmark steady-state. In this case:

$$\tilde{C}_T = r\tilde{F} + \tilde{Y}_N - \tilde{P}_T^X \tag{12}$$

where $\tilde{F} \equiv \mathrm{d}F/C_{T,0} = (1/\gamma)(\mathrm{d}F/Y_0)$. According to equation (12), consumption of traded goods is driven by the net foreign asset position, the level of tradable output endowment and the export prices. Log-linearizing around the steady state for the demand and supply of non-traded goods yields, respectively:

$$\tilde{Y}_N = \tilde{C}_N = \tilde{C}_T - q\,(\tilde{P}_N - \tilde{P}_T) \tag{13}$$

$$\tilde{Y}_N = \tilde{C}_N = \left(\frac{s-q}{s-1}\right) g\,(\tilde{P}_N - \tilde{P}_T) + \left(\frac{s}{s+1}\right) \tilde{A}_N \tag{14}$$

Note that equation (14) includes the impact of productivity surges in non-tradables, $\tilde{A}_N$. Combining and rearranging equations (12)-(14), we find the expression for the relative price of non-traded goods:

$$\tilde{P}_N - \tilde{P}_T = \frac{1+s}{q(1+s)+g(s-q)} \left[r\tilde{F} + \tilde{Y}_T + \tilde{P}_T^X - \frac{s}{1+s}\tilde{A}_N \right] \tag{15}$$

with its foreign counterpart defined analogously,[7]

$$\tilde{P}_N^* - \tilde{P}_T^* = \frac{1+s}{q(1+s)+g(s-q)} \left[-\left(\frac{n}{1-n}\right) r\tilde{F} + \tilde{Y}_T^* + \tilde{P}_T^M - \frac{s}{1+s}\tilde{A}_N^* \right] \tag{16}$$

2.3. The Real Exchange Rate Equation

In equation (5) we defined the real exchange rate as the ratio of domestic to foreign consumer price index, that is, $Q = P/P^*$. Under the assumption of similar preferences, we plug eq.(4) and its foreign counterpart into eq.(5) to define the log of the real exchange rate,

$$q_t = p_t - p_t^* = \gamma(p_{Tt} - p_{Tt}^*) + (1-\gamma)(p_{Nt} - p_{Nt}^*) \tag{17}$$

where lowercase letters indicate the natural log of uppercase letters, *i.e.* $x=\ln X$. Rearranging terms in (17), we obtain the Engel (2000) decomposition of the real exchange rate:

[7] We assume that the tradable output that is exported from the Foreign country is entirely consumed by the Home country in a two-country world. Hence, $P_T^{*X} = P_T^M$.

$$q_t = x_t + y_t = (p_{Tt} - p_{Tt}^*) + (1-\gamma)(p_{Nt} - p_{Tt}) - (1-\gamma)(p_{Nt}^* - p_{Tt}^*) \qquad (18)$$

where $x_t \equiv p_{Tt} - p_{Tt}^*$ denotes the relative price of traded goods, and the relative price of non-traded to traded goods is $y_t \equiv (1-\gamma)(p_{Nt} - p_{Tt}) - (1-\gamma)(p_{Nt}^* - p_{Tt}^*)$. According to Engel (2000), x_t is expected to be a stationary process. Although large and persistent, deviations from the law of one price in traded goods are stationary (Engel, 1993; Wei and Parsley, 1995). This result holds even in the presence of transportation costs (Obstfeld and Taylor, 1997). In this case, the unit root behavior in real exchange rates (q_t) might be induced by non-stationary behavior of y_t. The non-stationarity of y_t could be driven by permanent technology shocks, permanent demand shocks or permanent terms of trade shocks. From equation (18), we obtain the changes in real exchange rate:

$$\tilde{q}_t = \tilde{x}_t + \tilde{y}_t = (\tilde{P}_{Tt} - \tilde{P}_{Tt}^*) + (1-g)(\tilde{P}_{Nt} - \tilde{P}_{Tt}) - (1-g)(\tilde{P}_{Nt}^* - \tilde{P}_{Tt}^*) \qquad (19)$$

Thus, plugging (15) and (16) into (19) yields the equation for the real exchange rate changes.

3. Data and Empirical Implementation

3.1. The Data

In order to test the long-run equilibrium relationship between the real exchange rate and its determinants, we pool a sample of 67 countries across the world over the 1966-97 period. In what follows, we present the definition and sources of the data used in our empirical evaluation.

First, our dependent variable, the real exchange rate (q), is the CPI-based multilateral real exchange rate. It was constructed as the ratio between the domestic price index (converted in dollar terms at the period average nominal exchange rate) and a trade-weighted average of the trading partners' price indices also expressed in US dollar terms, according to the following formula:

$$q = \frac{(P/e)}{\prod_k (P_k / e_k)^{d_k}}$$

where P is the consumer price index (CPI) of the domestic country, e represents the exchange rate (price of the US dollar in units of local currency), P_k and e_k represent the CPI and exchange rate for the trading partners, and δ_k represent the *IMF*-generated weights based on both bilateral trade shares and export similarity. According to this definition, an increase in the real exchange rate implies a real appreciation of the domestic currency.[8]

[8]Lane and Milesi-Ferreti (2000) argue that cross-country differences in the construction and coverage of WPI indices, together with their more limited availability, implies that the CPI based real exchange rate measures are more reliable. In addition, the existence of sizable black market premium is an important

The net foreign assets data (*NFA*) is primarily drawn from Kraay, Loayza, Servén and Ventura (2000). This database comprises a set of foreign asset and liability stocks for a large group of industrial and developing countries spanning over the 1960-97 period.[9] The construction of the data is thoroughly documented in Kraay *et al.* (2000), and we define the net foreign asset position (NFA) of country *i* in year *t* as:

$$NFA_{it} = [FDIA_{it}\text{-}FDIL_{it}] + [EQYA_{it}\text{-}EQYL_{it}] + [RA_{it} + LA_{it} - LL_{it}]$$

where the letters *A* and *L* denote assets and liabilities, respectively. Thus, the net foreign asset position is the sum of net holdings of direct foreign investment, FDIA-FDIL, plus net holdings of portfolio equity assets, EQYA-EQYL, and the net position in non-equity related assets (*i.e.* "loan assets"). In turn, the net position in non-equity related assets consists of international reserves, RA, and the net loan position, LA-LL. Finally, for the discussion of the main issues related to the construction of the data, see Appendix A.

Regarding the relative productivity of the traded sector, we compute the ratio of output per worker for country *i* relative to the rest of the world. To perform this task we collect data on GDP per worker for country *i* as well as its trading partners from the Summers-Heston data on GDP per capita updated to 1997 by the World Bank. To compute the foreign productivity, we follow the same methodology and trade weights used for the construction of the real exchange rate. On the other hand, non-traded productivity is measured by the value added per capita for the construction sector, which is deflated using the PPP investment prices. The foreign productivity in this sector is computed following the same methodology and trade weights used for the construction of the real exchange rate. Finally, the data on the terms of trade, P_T^X / P_T^M, is standard and, thus, obtained directly from the World Bank's World Development Indicators (WDI) database.

3.2. Empirical Implementation of the Model

We already defined the log of the real exchange rate for country *i* at time *t*, as the sum of the relative price of traded goods (x_{it}) and the relative price of non-traded to traded goods (y_{it}):

$$\begin{aligned} q_{it} &= x_{it} + y_{it} \\ x_{it} &= (p_{T,it} - p_{T,it}^{*}) \\ y_{it} &= (1-\gamma)(p_{N,it} - p_{T,it}) - (1-\gamma)(p_{N,it}^{*} - p_{T,it}^{*}) \end{aligned} \tag{20}$$

Recent theoretical and empirical evidence shows that the relative price of traded goods might be bounded below and above by transaction costs. Specifically, deviations from the

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issue for developing countries in the 70s and 80s, which prevents the use of official exchange rates in those years.

[9]The database excludes "small island economies" (specifically, those with population under 1 million in 1995) as well as former socialist economies. Small economies were excluded because they tend to display higher volatility than larger economies (Easterly and Kraay, 1999), and this would add too much noise to our empirical experiments. In addition, they also include a number of tax havens attracting disproportionately large financial flows, which would distort the cross-country dimension of the data. On the other hand, former socialist economies were excluded because data availability was too limited.

law of one price follow a non-linear process in spatially separated market (Sercu and Uppal, 2000). This implies that x_{it} is a mean reverting process with the speed of adjustment varying directly with the extent of the deviation from PPP.

In practice, we assume that deviations from the law of one prices in traded goods depends on transaction costs. Given the lack of cross-country data on transaction costs and relying on the empirically successful gravity equation model (Leamer and Levihnson, 1995; Engel and Rogers, 1996), we assume that these costs depend on geographical factors. Furthermore, transaction costs (as well as tariffs) have been decreasing over time due to world trade agreements (i.e. global shocks to tradable prices). Therefore, deviations from the law of one price in tradables might be approximated by country-fixed effects, that is:

$$\tilde{x}_{it} \equiv \tilde{p}_{T,it} - \tilde{p}^{*}_{T,it} = f(i) + g(t)$$

or,

$$x_{it} = \boldsymbol{h}_i + \boldsymbol{m}_t \qquad (21)$$

where $f(i)$ and $g(t)$ are functions that depend on the cross-sectional and time series dimensions, respectively. Hence, $\boldsymbol{h}_i$ captures the country-specific fixed effects associated to the gravity equation, whereas $\boldsymbol{m}_t$ captures the global shock of trade policies on the relative price of tradables. On the other hand, we know that changes in the relative price of non-traded to traded goods can be found by plugging (15) and (16) into $\tilde{y}_{it}$:

$$\tilde{y}_{it} = \left(\frac{\Psi}{1-n}\right) r\tilde{F}_{it} + \Psi(\tilde{Y}_T - \tilde{Y}^{*}_T)_{it} + \Psi(\tilde{P}^X_T - \tilde{P}^M_T)_{it} - \left(\frac{\boldsymbol{s}\Psi}{1+\boldsymbol{s}}\right)(\tilde{A}_N - \tilde{A}^{*}_N)_{it} + \boldsymbol{x}_{it} \qquad (22)$$

where $\Psi = (1-\gamma)(1+\sigma) / [\theta(1+\sigma)+\gamma(\sigma-\theta)]$, $0<\Psi<1$ if $\sigma>\theta$, and ξ_{it} represents any deviation from the first order conditions (7)-(9), which we expect to be stationary. Note that eq.(22) can also be expressed in log levels:

$$y_{it} = \frac{\Psi r}{(1-n)\boldsymbol{g}}\left(\frac{F}{Y}\right)_{it} + \Psi \ln\left(\frac{Y_T}{Y^{*}_T}\right)_{it} + \Psi \ln\left(\frac{P^X_T}{P^M_T}\right)_{it} - \left(\frac{\boldsymbol{s}\Psi}{1+\boldsymbol{s}}\right)\ln\left(\frac{A_N}{A^{*}_N}\right)_{it} + \boldsymbol{x}_{it} \qquad (23)$$

Finally, the equilibrium real exchange rate equation (i.e. the fundamental equation of our empirical framework) is obtained by plugging equations (21) and (23) into (18):

$$y_{it} = x_{it} + y_{it} = \boldsymbol{h}_i + \boldsymbol{m}_t + \boldsymbol{b}_1\left(\frac{F}{Y}\right)_{it} + \boldsymbol{b}_2 \ln\left(\frac{Y_T}{Y^{*}_T}\right)_{it} + \boldsymbol{b}_3 \ln\left(\frac{P^X_T}{P^M_T}\right)_{it} + \boldsymbol{b}_4 \ln\left(\frac{A_N}{A^{*}_N}\right)_{it} + \boldsymbol{x}_{it} \qquad (24)$$

Among the main predictions of equation (24), we have the following: First, we expect that countries with significant external liabilities need to run trade surpluses in order to service them, and for this reason they require a real exchange rate depreciation (*"transfer effect"*). On the other hand, Obstfeld and Rogoff (1995) claim that a transfer from Home to Foreign country reduces the domestic wealth and hence raises labor supply and the supply of exportables, thus affecting the relative price (we expect that $\beta_1>0$). Second, the relative price of non-traded goods must be growing faster at home than abroad if the ratio of traded to non-traded goods productivity is growing faster at home than abroad. Furthermore, if we assume that the price of tradables equalize, the price of home national output must be rising relative to the price of foreign national output. Hence, if traded goods productivity relative to non-traded goods productivity is growing faster at home than abroad, home currency should appreciate in real terms (*i.e.* Balassa-Samuelson effect). Hence, we expect that $\beta_2>0$ and $\beta_4<0$. Finally, terms of trade improvements would increase the consumption of tradables and generate positive wealth effects that reduce the labor supply to the non-traded sector. This leads to an increase in the relative price of non-tradables and hence an appreciation of the real exchange rate (we expect that $\beta_3>0$).

4. Results

We estimate the real exchange rate equation specified in (24) using recently developed panel cointegration. First, we test the presence of unit roots in the series involved in our analysis. Second, we test for cointegration and we present our results. Given the heterogeneity of our sample, we also estimate the long-run real exchange rate equations for sub-samples of countries classified by income per capita and capital controls. Finally, we explore the short-run dynamics by characterizing the half-life of real exchange rate deviations from the equilibrium.

4.1. Testing for Unit Roots

Before estimating the long-run real exchange rate equation, we test for unit roots on all the series involved in our analysis (*i.e.* real exchange rates, relative output, the net foreign asset position, the terms of trade, and the relative productivity of the non-traded sector). Instead of applying the low power country-by-country unit root tests, we follow the strategy followed by Im, Pesaran, and Smith (1995), who developed a panel unit root test for the joint null hypothesis that every time series in the panel is non-stationary. Im, Pesaran and Smith (IPS) propose a testing procedure which averages all individual unit root test statistics.[10] The basic regression framework is the following:

$$y_{it} = \rho_i y_{i,t-1} + \sum_{k=1}^{p_i} \phi_{ik} \Delta y_{i,t-k} + z_{it}'\Gamma + \xi_{it} \tag{25}$$

with the null hypothesis of non-stationarity (H_0: $\rho_i=1$, for all i) and the alternative being H_1: $\rho_i<1$, for some i.[11]

[10]If the data from each country are statistically independent then, under the null, we can regard the average t-value as the average of independent random draws from a distribution with known expected value and variance (that is, those for a non-stationary series). This provides a much more powerful test of the unit root hypothesis than the usual single time series test.

[11]An analysis of the power of IPS test against alternatives is key for empirical work. If the test has high power, a few stationary series might drive the rejection of the unit root null and mislead us to model the panel as stationary. On the other hand, with low power tests, we might conclude that the panel contains a common unit root even if a majority of the series is stationary. In this context, Karlsson and Lothgren (2000) find that the power of the IPS test increases monotonically with: (i) a higher number N of series in the panel; (ii) a larger time dimension T in each individual series, and (iii) a higher proportion of stationary series in the panel.

Im, Pesaran and Smith compute a $\bar{t}$-statistic, which is an average of the individual ADF statistics,

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t(\rho_i)$$

where $t(\rho_i)$ is the individual t-statistic of testing the null hypothesis in equation (25). The critical values are tabulated by Im, Pesaran and Smith (1995). In general, panel unit root tests suffer from a dramatic loss of power if individual specific trends are included. This is due to the bias correction that also removes the mean under the sequence of local alternatives. Simulation results indicate that the power of IPS tests is very sensitive to the specification of deterministic trends (Breitung, 1999).[12]

Recently, Taylor and Sarno (1998) have proposed a multivariate tests in which the null hypothesis is that at least one of the series in the panel is a realization of a unit root process. This null hypothesis is violated if all of the series are in fact realizations of stationary processes (*i.e.* $H_1 : \rho_i < 1$, for all *i*). The test procedure is a special application of Johansen's (1988) maximum likelihood procedure for testing for the number of cointegrating vectors in a system. Although this alternative hypothesis is more rigorous towards stationarity than the one presented by IPS, we can not perform this test because we lack the sufficient time dimension *T*, given the number *N* of countries involved.

In order to implement panel unit root tests, we need to remove any common time-effects prior to carrying out the augmented Dickey-Fuller (ADF) regressions. Hence, we regress the variable on a set of time dummies and take the residuals, thus, reducing the risk of correlation across countries. Unit root tests for the series in levels include a time trend and five (5) augmenting lags. In addition, we test for stationary in first differences in order to check if our panel series are I(1). In this case, the alternative implies stationarity around a constant since any time trend levels will be removed by differencing.

Panel unit root tests are performed over the largest time-series sample available for all countries. We have data for the real exchange rate ($\ln q$) and the relative GDP per worker ($\ln Y_T/Y_T^*$) over the 1960-97 period. However, complete data for series such as the terms of trade ($\ln P_T^X/P_T^M$), the ratio of net foreign assets to GDP (F/Y) and the relative non-traded productivity ($\ln A_N/A_N^*$) is available over the 1966-97 period. We report the results in Table 1. There, we fail to reject the null hypothesis for the series in levels, but not for the series in first differences. Hence, we find evidence that all the series involved in our analysis are non-stationary in levels and in every case we reject a unit root in first differences (i.e. stationary in differences). That is, all our series are integrated of order one, *i.e.* I(1).

[12] Choi (1999) has demonstrated that the empirical size of the IPS test is reasonably close to its nominal size *0.05* when *N* is small, and that is has the most stable size among the panel unit root tests. However, when a linear time trend is included in the model, the power of all tests (including IPS) decrease considerably.

4.2. Testing for Cointegration

The estimation results that we present in Tables 4 through 7 would be misleading if we do not test for cointegration in our panel data model.[13] We employed cointegration tests proposed by Kao (1999), Pedroni (1995), and Pedroni (1999) to test whether the cointegration relationship exists in the estimated real exchange rate equation.

Kao (1999) has computed two types of panel cointegration tests, four DF-type tests, DF_ρ, DF_t, DF_ρ^*, DF_t^*, and an ADF-type test. Regarding the DF-type tests, the first two are based on the strong exogeneity of the regressors and errors, whereas the last two test for cointegration in the presence of endogenous regressors. Finally, Kao found that the distribution of these statistics converges to a standard normal distribution.[14]

Pedroni (1995) developed two sets of cointegration tests in panel data models allowing for considerable heterogeneity. The first set of statistics involved averaging test statistics for cointegration in the time series across individuals, which included an average of the Phillips and Ouliaris (1990) statistic. On the other hand, the second set performs the averaging by pieces, so that the limiting distributions are based on limits of piece-wise numerator and denominator terms. This set included four panel variance ratio statistics. According to these tests, the rejection of the null hypothesis means that enough of the individual cross-sections have statistics "far away" from the theoretically predicted means if they were generated under the null.[15]

In Table 2, we present the results of our cointegration tests. We find that all test statistics are significant so that the null of no cointegration is strongly rejected. Hence, there is evidence in support of the cointegration hypothesis.

4.3. Correlation Analysis

In order to have a first approximation to the comovement between the real exchange rate and its determinants, we perform a correlation analysis. Given that our variables are integrated of order one, we compute the correlation between the average of the log differences in the real exchange rate and its determinants. This analysis is conducted for a cross-section of 67 countries over the 1966-97 period and for a panel data of non-overlapping 8-year observations over the same period. Taking advantage of the panel

[13] Note that given the results in Kao (1999), our coefficient estimates will be consistent whether or not we have a cointegrating relationship. The estimates would be reliable provided the model being estimated represents the true relationship generating the data. An omitted variable will produce a bias in our estimates. The key question is really whether the error term is independent noise or the product of miss-specification.

[14] For a summary of recent developments in panel unit root and cointegration tests, see Baltagi and Kao (2000), Kao, Chiang and Chen (1999) and Pedroni (1999).

[15] Pedroni (1999) derives asymptotic distributions and critical values for several residual based tests of the null of no cointegration in panels when there are multiple regressors, with considerable heterogeneity being allowed across individual members of the panel regarding to the associated cointegrating vectors and the dynamics of the underlying error process. Results with these test statistics support the hypothesis of cointegration. Although not reported, they are available from the author upon request.

data, we also compute the evolution of the correlation across sub-periods. Finally, we compute cross-section and panel data correlations for different subgroups of countries. From the results reported in Table 3, we find the following:

First, the real exchange rate , *Dln q*, and net foreign assets, *D(F/Y}*, are positively and significantly correlated for both the cross-section and panel data sets. However, the sign of the correlation is not robust across sub-periods (*i.e.* it is only positive in 1990-97) or across subgroups of countries. We find a higher positive association for countries with low capital controls in the cross section (*0.19*), and for high and upper-middle income countries in the panel data set (*0.06*).

Second, changes in productivity, $Dln(y_T/y_T^*)$, are positively and significantly correlated with *Dln q*. The sign and significance level of this correlation is robust across the different groups of countries for both the cross-section and panel data sets. Although we find that cross-section correlation are stronger in low-income countries than in high-income countries (0.33 vs. 0.28), we find the opposite result in panel data (0.18 vs. 0.42, respectively). In addition, we find that the cross-section correlation between $Dln(y_T/y_T^*)$, and *Dln q* is stronger in the sample of countries with low capital controls than in the sample of countries with high capital controls (0.43 vs. 0.17), whereas it is only slightly larger for the panel correlations (0.25 vs. 0.24, respectively).

Third, terms of trade shocks, $Dln(P_T^X/P_T^M)$, and *Dln q* are positively associated. The cross-section correlation is significant only for the sample of all countries (0.32), low income countries (0.45) and countries with low capital controls (0.47). However, panel correlations are positive and significant for all the subgroups of countries. This correlation seems to be slightly larger in high income than in low-income countries (0.26 vs. 0.24), and it is larger in countries with high capital controls than in countries with low capital controls (0.27 vs. 0.20).

Finally, contrary to what we expected, we find a positive association between changes in non-traded productivity , $Dln(A_N/A_N^*)$, and *Dln q*. This result holds across sub-groups of countries for both cross-section and panel data sets.[16] The finding of a positive unconditional correlation between $Dln(A_N/A_N^*)$ and *Dln q* might be attributed to the presence of both demand and technology shocks with offsetting forces that affect this association. For this reason, we compute the condition correlation between these two variables, controlling for demand shocks (*i.e.* using private consumption growth as a control). We find that controlling for demand shocks, the correlation between $Dln(A_N/A_N^*)$ and *Dln q* is negative for both the cross-section and panel correlation across countries, though statistically significant only in the latter.

[16] However, this correlation is negative (as expected) only for the sample of all countries in 1974-81 and 1990-97.

4.4. Panel Cointegration: The Evidence

The main goal of the present paper is to estimate the equilibrium path of the real exchange rate using a sample of 67 countries across the world over the 1966-97 period. To perform this task, we estimate the long-run relationship specified in equation (24) using panel cointegration techniques. Recently, there has been an extensive research on panel cointegration (Kao, 1999; Pedroni, 1999; Phillips and Moon, 1999; Kao and Chiang, 2000).

4.4.1 On the Estimation of Panel Cointegrated Models: A Very Brief Review

Among the most used estimators in panel cointegrated models we have the OLS, fully-modified least squares (FM), and dynamic least squares (DOLS) estimators. Kao and Chiang (2000) have derived the limiting distributions for the FM and DOLS estimators in a cointegrated regression and showed that they are asymptotically normal. Phillips and Moon (1999) also obtained similar results for the FM estimator. However, Kao and Chiang (2000) find that in finite samples: (i) the OLS estimator has a non-negligible bias in finite samples, (ii) the FM estimator does not improve over the OLS estimator in general, and (iii) the DOLS estimator may be more promising than OLS or FM estimators in estimating the cointegrated panel regressions.[17]

Consider the following panel regression with fixed effects,

$$y_{it} = \alpha_i + x_{it}'\beta + u_{it} \ ; \ i=1,...,N; t=1,...,T \tag{26}$$

where y_{it} represents the real effective exchange rate (log q_{it}), β is a $4x1$ vector of slope parameters, α_i represents the intercepts, u_{it} are the stationary disturbance terms, and $\{x_{it}\}$={F/Y, ln(Y_T/Y_T^*), ln(P_T^X/P_T^M), ln (A_N/A_N^*)} represent the vector of real exchange rate determinants. We assume that $\{x_{it}\}$~I(1), for all i,

$$x_{it} = x_{i,t-1} + \varepsilon_{it} \tag{27}$$

Equations (26) and (27) specify a system of cointegrated regressions also known as *"triangular representation"* (Phillips, 1991). Given the superiority of the DOLS estimator among the alternative techniques (Kao and Chiang, 1999), we decide to apply DOLS to our real exchange rate equation.[18] The DOLS estimator, $\hat{\beta}_{DOLS}$, can be obtained by running the following regression:

$$y_{it} = \alpha_i + x_{it}'\beta + \sum_{j=-k_1}^{k_2} c_{ij}\Delta x_{i,t+j} + v_{it} \tag{28}$$

For details on the limiting distribution of the OLS and DOLS estimators, see Appendix B.

[17]The superiority of DOLS estimators over FM and OLS is practically demonstrated in Kao, Chiang and Chen's (1999) evaluation of international R&D spillover regressions.

[18]The DOLS estimator is a panel version of the dynamic least squares estimator proposed by Phillip and Loretan (1991), Saikkonen (1991), and Stock and Watson (1993).

4.4.2 Estimation Results

We pool our sample of 67 countries according to different sub-sample of countries and time periods and present the panel evidence in Tables 4 through 7. Note that in Tables 1 and 2 we have already shown the non-stationary of our series in levels as well as the stationarity of the residuals in the cointegrating relationship, respectively. This evidence suggests that our real exchange rate equation, as specified in (24), is valid in the long-run. As we explained above, our preferred estimation technique is the dynamic least squares (DOLS) with 2 lags and 1 lead, that is, DOLS(2,1).[19]

A. Full Sample of Countries

In Table 4 we report the coefficient estimates for the full sample of countries (67), using: (i) different estimation techniques (OLS, OLS with bias correction, FM, and DOLS); (ii) different indicators of productivity in the traded sector (output per worker *vs*. TFP); and, (iii) different sample periods (1966-97 *vs*. 1973-97).

OLS estimates reported in Table 4 have the expected signs (for 1966-97) and their computed t-statistics are large. However, these estimates are generally biased due to endogenous regressors. We also present estimates based upon OLS with bias correction, FM and DOLS, respectively.[20] The DOLS estimated coefficients are quite different from the FM estimator, even though both estimators have the same limiting distribution (Kao and Chiang, 1999).[21]

First, we find that regardless of the estimation technique used, our coefficient estimates for the 1966-97 period have the expected signs and that they are statistically significant (except for the coefficient of net foreign assets for the OLS estimates). Second, we find that our OLS estimates are downward biased. If we compare our DOLS and OLS estimates, we clearly find that the former is significantly larger than the latter ones. Third, both FM and DOLS estimates have the expected signs regardless of the productivity measure and the sample period chosen. Finally, it seems very restrictive to assume *"country homogeneity"* in our estimates for the long-run real exchange rate equation. The transmission channels could differ among countries due to different levels of income or the imposition of capital controls. For this reason, we proceed to estimate eq.(24) for different sub-samples of countries. Table 5 presents coefficient estimates for sample of high and upper-middle income countries as well as for low and lower-middle income countries. In Tables 6 and 7 we present estimates for subsample of countries according to the presence of capital controls and the degree of black market premium in the foreign

[19] Recall that we modeled the impact of transaction cots in international trade using geographical factors in the line of the gravity model. This geographical factors are captured in our panel data model as the country fixed-effect η_i in equation (24).

[20] The FM estimation corrects the dependent variable using the long-run covariance matrices for the purpose of removing the nuisance parameters and applies the usual OLS estimation method to the corrected variables.

[21] Note that the DOLS estimator includes lead and lag terms to correct the nuisance parameter in order to obtain coefficient estimates with nice limiting distribution properties.

exchange market, respectively.[22] In addition, we also report estimates changing the sample time period to 1973-97 in order to check the robustness of our estimates for the post-1973 oil-shock (which is also a period dominated by floating rates in the industrial economies).

B. Sub-Samples according to Income Levels and Capital Controls

As we stated before, we first tested the robustness of our long-run estimates by adjusting the real exchange rate equation in different sub-samples according to income levels and capital controls. In the former case, we select the criterion used by the World Bank to classify the countries by income level[23], and we divide our sample into high and upper-middle income countries (33) and low and lower-middle income countries (34). Regarding the latter case, we used two different definitions. First, we defined the subsample of countries according to the presence of capital controls by computing the sum of capital control dummies (1 for the presence of the restriction, and 0 otherwise) collected from the IMF's Exchange Arrangements and Exchange Restrictions. These dummies capture the presence of: (a) multiple exchange rate practices, (b) current account restrictions, (c) capital account restrictions, and (d) surrender of exports proceeds. If the sum of these four categories was higher than or equal to three (i.e. presence of at least three restrictions) over the 1966-97 period, we consider it a country with high capital controls. Second, we consider a country with high capital controls if the average black market premium (BMP) over the 1966-97 period was higher than 20 percent. Among the main results, we have:

First, we find a positive and significant relationship between the real exchange rate and the net foreign assets for the full sample of countries. This result also holds for both the sample of high-income countries and low income countries, with the coefficient of low income countries being larger (0.36 vs. 0.18). This denotes a more powerful transfer effect for the low and lower-middle income countries. On the other hand, we find a positive coefficient for the net foreign assets only for the sub-samples of countries with low capital controls (0.16) and countries with low black market premium (0.19).[24]

Second, we find a positive and significant long-run impact of relative productivity on the real exchange rate. This result is robust whether we use different indicators of relative productivity (that is, relative output per worker or relative total factor productivity) or

[22]For a complete list of countries and its classification according to income level or capital controls, see the Table in the appendix of results.

[23]Using the GNP per capita, the World Bank classifies economies as low income, middle income (subdivided into lower middle and upper middle), or high income. Low-income and lower-middle income economies are sometimes referred to as developing economies. According to the latest World Development Report (The World Bank, 2000), economies are divided among income groups according to 1999 GNP per capita. The groups are as follows: low income, US$ 755 or less; lower-middle, US$ 756-2995; upper middle income, US$ 2996-9265; and high income, US$ 9266 or more.

[24]According to regression 1 in Table 5 (i.e. sample of high and upper-middle countries over the 1966-97 period), if the net foreign asset position of Norway (3.9 percent of GDP) improves in such a way that reaches the position of the Germany (10.9 percent of GDP), the real exchange rate will experience an appreciation of 1.2 percent. On the other hand if Korea (-10.3 percent of GDP) reaches the asset position of Japan (10.4 percent of GDP), the real exchange rate appreciates by 3.7 percent.

different time periods. This result not only holds for the full sample but also across all the sub-samples presented from tables 6 through 8. It is interesting to note that the coefficient of relative productivity, $\ln(Y_T/Y_T^*)$, is high in the sample of high-income countries than in the sample of developing countries.[25] Consistent with this result, we find that the impact of $\ln(Y_T/Y_T^*)$ is higher in the sample of countries with lower capital controls and low black market premium. Using regression 1 of Table 4 (i.e. sample of all countries over the 1966-97 period), an increase in relative productivity of 1 percent might generate a real appreciation of 1.3 percent in the exchange rate.

Third, we find that a decline in the terms of trade might be associated with a real exchange rate depreciation. This positive relationship between the terms of trade and the real exchange rate holds for the full sample and for the sample of both industrial and developing countries. Interestingly enough, this positive and significant association holds for high-income countries but not for low-income countries, whereas the relationship is not clear if we classify the countries according to capital controls. According to our estimates in regression 1 of Table 4, a one percent decline in the terms of trade might be associated with a 0.7 percent depreciation in the real exchange rate.

Fourth, we find that surges in non-traded productivity might depreciate the real exchange rate. We find that the relative non-traded productivity, $\ln(A_N/A_N^*)$, has a negative and significant coefficient for the full sample of countries as well as for the sample of industrial countries. The coefficient is negative and significant only for the sample of high and upper-middle income countries (see Table 5). In addition, this finding holds for the samples of countries with low capital controls and low black market premium, with the long-run impact being larger for the former sample. Using the estimates of regression 1 in Table 4, we find that a 1 percent surge in productivity in the non-traded sector is associated with a *0.2* percent depreciation in the real exchange rate.

5. Robustness Checks on the Long-Run Coefficient Estimates

In the present section we test whether our long-run estimates are valid for all countries and/or groups of countries as well as over time. In addition, we evaluate the consistency of our estimates by comparing them to their calibrated theoretical values.

5.1. Testing for Group and Country Heterogeneity

In Tables 8 through 10 we present formal tests of homogeneity between and within groups of countries. First, we test the equality of the coefficient estimates of the long-run real exchange rate equation between different subgroups of countries classified according to income levels and capital controls (see bottom panels of Tables 5-7). Second, we test the null of homogeneity across countries by formulating a Hausman-type test in the spirit of Pesaran, Shin and Smith (1999). In this case, we compare our pooled DOLS estimates, already presented in the previous sections, with the average of the DOLS estimates performed on a country-by-country basis. After computing these two estimators, we will

[25] An analogous result holds when we compare industrial economies and developing countries.

test the null that these two coefficient estimates are equal by formulating both individual and joint Hausman tests.[26] See Tables 8-10 for more details.

A. Testing the Equality of Coefficients across Sub-Groups of Countries

First, we find that the long-run coefficient estimates for the high and upper-middle income countries are jointly statistically different from the ones for the low and lower-middle income countries. Individually, all coefficients but the ones for the net foreign assets (F/Y) are statistically different between high and low income countries (see bottom panel of Table 5).

Second, regardless of the criterion used to classify countries according to the use of capital controls, we find that the long-run coefficients of the real exchange rate equation for the sample of countries with low capital controls are jointly different from the coefficients of countries with high capital controls. Analogously, we find that all individual coefficients except for the one for net foreign assets, F/Y, are statistically different across groups. See bottom panel of Tables 6 and 7.

B. Testing the Hypothesis of Country Homogeneity

In this section we perform hypothesis testing of country homogeneity behind the long-run estimates presented in Tables 4-7. Here we test whether the *'pooling assumption'* is valid for our long-run estimates in the sample of all countries (Table 8) and in all sub-samples (Tables 9 and 10).

First, we test for the presence of country heterogeneity in the full sample of countries (see Table 8). At the 5 percent significant level, we find that the pooled and average DOLS estimates of F/Y and $\ln(A_N/A_N^*)$ are statistically equal for the 1966-97 period. The opposite result (*i.e.* country heterogeneity) holds for the coefficients of $\ln(P_X/P_M)$ and $\ln(y_T/y_T^*)$. On the other hand, the joint Hausman test rejects the null of homogeneity in the limit (p-value=0.059). Furthermore, if we evaluate the estimates for the 1973-97 period, we surprisingly find that although the individual tests reject the null of heterogeneity at the 5 percent level of significance, the joint test does not (p-value=0.073 for the joint test).

Second, we test the hypothesis of country heterogeneity for both sub-samples of high and upper-middle income countries and for low and lower-middle income countries (see Table 9). We find that the null hypothesis of country homogeneity is supported by both individual and joint Hausman tests for the sample of high-income countries. In contrast, individual and joint tests, except for the individual test for F/Y, favored the hypothesis of heterogeneity.

[26] In the case of the average country-by-country DOLS, we estimate separately the long-run real exchange rate equation for each country and we examine the distribution of the estimated coefficients across countries. According to Pesaran and Smith (1995), the mean of the country estimates will produce consistent estimates of the average of the parameters. However, this estimator does not take into account the fact that certain parameters may be similar across groups.

Third, we test the pooling assumption for both the sub-samples of countries with low capital controls and high capital controls (see Table 10). In this case, we find that both individual and joint Hausman tests fail to reject the null of homogeneity for the sample of low capital controls at the 5 percent significance level. The opposite result is found for the sample of countries with high capital controls, where we reject the null of homogeneity. Note that the results hold regardless of the criterion used for the capital control classification of countries.

In summary, we find that the notion of a long run equilibrium relationship similar across countries (i.e. the pooling assumption) is valid for the sub-samples of high and upper-middle income countries as well as for the sub-sample of countries with low capital controls.

5.2. Testing for Structural Change

Although there is a wide array of papers on structural change testing for non-stationary time series (Andrews, 1993; Campos, Ericsson, and Hendry, 1996), recent research has extended this analysis to non-stationary panel data. In order to assess the stability over time of the parameters in our real exchange rate equation, we apply the methodology developed by Kao and Chiang (2000) which we briefly explain in Appendix C.

Kao and Chiang (2000) propose a Wald-type test statistics for detecting breaks at unknown dates in panel data cointegrated regressions. They find that the limiting distribution of this Wald test is free of nuisance parameters, and similar but not identical to that of the test developed by Andrews (1993). That is, the limiting distribution is the square of a Bessel process. Finally, their test has non-trivial power against a wide array of alternatives (*i.e.* regardless of the particular type of structural change).

We find that there is a significant structural break for all samples of countries, although it is not similar across groups of countries. For the full sample, we find that there is a structural change in the coefficient estimates in 1973 (*i.e.* year of the oil crisis).[27] The same result (year 1973) is found for both the group of high and upper-middle income countries and countries with low-capital controls (see Table 11). However, we find that 1976 represents the year of structural change for low and lower-middle income countries, whereas 1985 is the year of structural break for the sample of countries with high capital controls.

5.3. Consistency Checks between estimates and calibrated values

In order to assess the plausibility of our coefficient estimates of the real exchange rate determinants, we compare our coefficient estimates with the calibrated parameters that we would obtain by using the parameter values in the existing literature on both industrial

[27] Although not reported in the Table, the value of the sum of squared residuals when considering 1985 as the year of structural change is very close to the sum of squared residuals when we adopt 1973 as the time break.

and developing countries. For our purposes we use the parameters calibrated by Stockman and Tesar (1995) for industrial countries. There, we find that the elasticity of intertemporal substitution (σ) is 0.5, the elasticity of intra-temporal substitution (θ) is 0.44, the share of tradables in the consumption basket (γ) is assumed to be 0.5, and the international real interest rate is 4 percent. On the other hand, for simplicity, we assume that developing countries would only differ in the elasticity of intertemporal substitution. Mendoza (1991) assume that this coefficient might take the value of 0.999. Finally, the country size of the representative industrial country is computed using the ratio of the (population-) weighted average of the GDP relative to the total GDP of the group of countries. Analogous computation is performed for the other group of countries. Using these calibrated parameters, we find the possible theoretical values for the coefficients of the real exchange rate determinants.

We find that the coefficient estimates for the full sample (see regressions 1 and 3 in Table 4) are broadly in line with the calibrated parameters. Using the sample of all countries over the 1966-97 period (regression 1), we find that the coefficient estimate for the net foreign asset position (F/Y) is larger than the calibrated parameter (0.2127 vs. 0.1494), whereas the estimated value for the relative productivity of the traded sector, $\ln(Y_T/Y_T^*)$, is quite similar to the calibrated value (1.3024 vs. 1.0870). On the other hand, the coefficient estimate for the terms of trade, $\ln(P_X/P_M)$, is smaller than the calibrated parameter (0.7427 vs. 1.0870), with the same result holding for the coefficient of the relative non-traded sector productivity, $\ln(A_N/A_N^*)$ (-0.1837 vs. -0.3623). Note that the latter coefficient estimate is similar to the calibrated parameter when we use the sample of all countries over the 1973-97 period (see regression 3 of Table 4). Now we reduce the elasticity of intertemporal substitution (σ) such that the coefficient estimate for the relative traded-sector productivity, $\log(Y_T/Y_T^*)$, and the calibrated parameter are equal. Based on the regression 1 of Table 5, we set σ=0.294, and hence, Ψ =1.3024, *i.e.* calibrated parameter for $\log(Y_T/Y_T^*)$ and $\log(P_X/P_M)$. Here, the estimate of F/Y is closer to the calibrated parameter (0.2127 vs. 0.1792, respectively). On the other hand, the coefficient of $\log(P_X/P_M)$ is not equal to the coefficient of $\log(Y_T/Y_T^*)$, being smaller than the calibrated parameter (0.7427 vs. 1.3037). Finally, the coefficient of $\log(A_N/A_N^*)$ is also smaller (in absolute value) than the calibrated value (-0.1856 vs. -0.2961).

On the other hand, the discrepancies between the coefficient estimates and calibrated parameters for the sample of high and upper-middle income countries are slightly larger than the ones obtained for the full sample of countries. Using the results from regression 1 of Table 5, we find that the estimated coefficient of F/Y doubles the value of the calibrated parameter (0.1804 vs. 0.0830). On the other hand, the coefficient of log (P_X/P_M) is slightly larger than the calibrated parameter (1.3649 vs. 1.0870), whereas the coefficient of $\log(A_N/A_N^*)$ is four times larger than the calibrated parameter (-1.4078 vs. -0.3623). In what follows, we modified the value of σ such that the estimated and calibrated coefficient for the terms of trade are equal (*i.e.* σ =0.2715 and Ψ =1.3649). We find that the estimated coefficients for the net foreign asset position and the non-traded sector productivity are closer to the calibrated parameters (0.1021 and -0.2857, respectively), however, we still have different coefficient estimates for the terms of trade and the traded sector productivity.

6. Investigating the Short-run Dynamics of Real Exchange Rates

After finding robust estimates for the long-run real exchange rate equation, we proceed to analyze its short-run dynamics. We constrained the coefficients in the long run relationship to be equal across countries, though we allow for different intercepts.[28] However, there is no reason to suppose that the speeds of adjustment or convergence to equilibrium would be the same. In fact, cross-country differences in the speed of adjustment towards equilibrium could be attributed to differences in trade policy, problems of information in the financial markets and transactions costs.

In practice, we take the cointegrating vector estimated for the full sample of countries over the 1966-97 period and we estimate the short-run relationship (i.e. error correction specification) country by country. For each country *i* at time *t*, we define the real exchange rate deviations from equilibrium as $d_{it} = q_{it} - x_{it}'\beta$. Hence, the error correction model for each country *i* can be specified as:

$$d_{it} - d_{i,t-1} = \phi_i \, d_{i,t-1} + \zeta_{it} \tag{29}$$

where ζ_{it} is an error term. If there is convergence, ϕ_i should be negative ($-1<\phi_i<0$), and the absolute value of ϕ_i should be interpreted as the annual decay for the real exchange rate deviations from the equilibrium, with the implied half-life of these deviations being $\ln(0.5)/\ln(1+\phi_i)$. The model specified in equation (29) is what we would call model ECM_1. On the other hand, we will estimate a broader specification to (29), which includes the changes in relative money supply. The rationale for including this variable is the impact of short-run money fluctuations on exchange rate volatility.[29] Hence, the alternative short-run specification (*i.e.* model ECM_2) is:

$$d_{it} - d_{i,t-1} = \phi_i \, d_{i,t-1} + \varphi_i \, \Delta(m-m^*)_{it} + \zeta_{it} \tag{30}$$

and given that $\Delta(m-m^*)$ is a significant variable in explaining real exchange rate volatility both theoretically and empirically, we describe our results using the estimates of equation (30).

We first estimate the speed of convergence for each country, $\hat{\phi}_i$ (*i=1,...,67*), and the implied half-lives. Then, we compute the median half-life of the deviations from the equilibrium for the full sample of countries as well as for sub-samples of countries classified by income per capita and capital controls. In general, we find that $\hat{\phi}_i$ is negative for all countries, which is consistent with the notion of convergence towards the steady state (*i.e.* existence of a cointegrating relationship). From our country estimates of

[28]We can assume a long-run equilibrium relationship among variables to be equal across countries due to budget or solvency constraints, arbitrage conditions or common factors influencing all groups in a similar fashion (Pesaran, Shin and Smith, 1999).
[29]We presume that by including the monetary variable, the speed of adjustment to the long run would be faster.

$\hat{\jmath}_i$, we find that the deviations from the equilibrium real exchange rate have an average annual decay rate of *21.2* percent, with an implied half-life of *3.3* years. In addition, given the heterogeneity between industrial and developing countries due to differences in trade policy, information in the financial markets or exchange rate regimes, we compute the implied half-life of real exchange rate deviations from the equilibrium for different sample of countries.

First, we find that deviations from the equilibrium real exchange rate decay at an average (median) rate of *24.2* (*24.8*) percent per year for our sample of high and upper-middle income countries, thus implying a half-life of *2.87* (*2.79*) years for these deviations. This result is close to the lower bound of the consensus interval of 2.5 to 5 years half lives of PPP deviations among studies using long-horizon data and panel data studies (Murray and Papell, 2002). On the other hand, the half-life of deviations from the equilibrium exchange rate is higher for the sample of low and lower-middle income countries (*3.8* years), with a rate of decay of *18.3* percent per year (see figure 1).

Second, we find that both the average and median rate of decay for countries with low capital controls are higher than the rates estimated for countries with high capital controls, thus implying that capital controls delay the convergence towards equilibrium. This result holds regardless of the criterion used to classify countries according to their use of capital controls (i.e. whether we use the dummy variable approach or the black market premium). Specifically, we find that deviations from equilibrium dissipate more rapidly in countries with low capital controls than in countries with high capital (i.e. a median decay rate of *22.3* and *15.5* percent, respectively). Hence, the implied half-life of real exchange rate deviations from equilibrium is shorter in countries with low capital controls than in countries with high capital controls (median half-lives of 3.1 and 4.5 years). See figures 1 and 2 for further details.

Finally, we attempt to roughly characterize the half-life of real exchange rate deviations from the equilibrium by evaluating their nexus with output per worker and capital controls. We find that the half-life of deviations from the equilibrium exchange rate are negatively associated with income per worker (*-0.13*), though the correlation is not significant. However, conditional to the accumulation of net foreign assets, this correlation is not only larger in absolute value (*-0.36*) but also statistically significant (see figure 3). On the other hand, we find that both the unconditional and conditional correlation between the half-life of real exchange rate deviations from the equilibrium and the capital controls (measured by either of the two proxies mentioned above) are negative and significant only in the case of the capital control index (see figure 4).

7. Conclusions

The lack of long time-series and suitable econometric procedures has been an obstacle to evaluating the long run real exchange rate behavior. However, we tackle these empirical problems by taking advantage of the recent panel data cointegration techniques. We use this econometric approach to estimate the real exchange rate equation derived from our theoretical model, which is in line with the new open economy macroeconomics. Given the heterogeneity of the sample, we also conduct the cointegration analysis for sub-samples of countries according to income per capita and the presence of capital controls. After estimating the long run coefficients, we attempt to characterize the short-run dynamics.

Among the most important findings in the present paper, we have:

- There exists a long-run relationship between the real exchange rates, the ratio of net foreign assets to GDP, the relative productivity of traded and non-traded sectors, and the terms of trade. This result holds regardless of the sample of countries and sub-periods analyzed.

- Consistent with Lane and Milesi-Ferreti (2000), we find evidence for the transfer effect in the full sample of countries, as well as in the sample of high and low income countries. However, we find that the impact of net foreign assets is stronger for the sample of low and lower-middle income countries.

- Relative productivity of the traded and non-traded sectors are associated with the real exchange rate following the pattern described by the *Balassa-Samuelson* hypothesis. Improvements in the ratio of Home to Foreign productivity in the traded sector might generate a real exchange rate appreciation, whereas a similar improvement in the non-traded sector productivity is associated to a real exchange rate depreciation. However, note that the impact of traded productivity is significantly larger than the impact of non-traded productivity surges.

- A secular decline in the terms of trade might lead to a real exchange rate depreciation. This result holds for the full sample of countries as well as for the sample of high-income countries.

- Robustness tests between and within sub-groups of countries reveal that the assumption of pooling the data to estimate a long-run equilibrium relationship similar across countries (i.e. test of country homogeneity) is valid for the samples of high and upper-middle countries and countries with low capital controls. In addition, joint Hausman tests reject the null of homogeneity for the full sample of countries in the limit (*p-value=0.059*).

- We find evidence of structural change in our real exchange rate equation. In general, we find that the first oil crisis (1973) represents a break in the long-run parameters for the full sample of countries as well as for the sample of high and upper middle income countries and countries with low capital controls. For the sample of low and lower-middle income countries, changes in parameter estimates occur in 1976.

- The estimated coefficients for the sample of all countries are in line with the calibrated parameters obtained from equation (26) using the empirical literature on Real Business Cycles. However, the discrepancy between the estimated and calibrated coefficient is larger in the case of high and upper middle income countries. Finally, note that if we decrease the elasticity of intertemporal substitution from *0.5* to approximately *0.3*, the discrepancy between the coefficient estimates and the calibrated parameters are smaller for both the full sample and the sample of high and upper-middle income countries.

- Deviations of the real exchange rate from the equilibrium are large and persistent, though there is robust evidence of convergence towards equilibrium. We find that deviations from the equilibrium decrease at an average annual rate of *21.2* percent, thus, implying a half-life of *3.3* years for these deviations. On the other, the implied half-life for deviations in high-income countries (*2.87* years) is consistent with the consensus interval of 2.5-5 years half-lives found in the literature (Murray and Papell, 2002). In addition, we find that half-life of the deviations from the equilibrium are negatively correlated with income levels, whereas the nexus with the degree of capital controls is not significant.

Finally, further research might attempt to extend our empirical implementation by including other shocks that generate deviations from long run PPP (e.g. government shocks). Also, the use of factor analysis would improve the evaluation of the real exchange rate deviations from the equilibrium by testing the existence as well as the importance of common factors explaining these deviations. Finally, a better understanding of the dynamic response of real exchange rates to different shocks in the economy (e.g. productivity shocks, demand shocks) could be achieved with the implementation of structural vector autoregression models.

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Appendix A: Methodological Issues on the Net Foreign Assets Data

The data on net foreign assets used in this paper is drawn from the recently developed database on Country Portfolios (Kraay, Loayza, Servén and Ventura, 2000). Although the construction of the data is thoroughly documented in Kraay *et al.* (2000), we will discuss some of the main issues in this appendix. We start by presenting the definition of the net foreign asset position (NFA) of country j in year t:

$$NFA_{jt} = NFQA_{jt} - NFLA_{jt} = [FDIA_{jt} - FDIL_{jt}] + [EQYA_{jt} - EQYL_{jt}] + [RA_{jt} + LA_{jt} - LL_{jt}] \tag{A.1}$$

where all variables are expressed in current US dollars. NFQA denotes the net holdings of equity-related assets and NFLA the net holdings of other assets. Using the letters *A* and *L* to denote respectively assets and liabilities, NFQA represents the net holdings of direct foreign investment, FDIA-FDIL, plus the net holdings of portfolio equity assets, EQYA-EQYL. In turn, NFLA represents the net position in non-equity-related assets, that for brevity we shall call "loan assets". The position consists of international reserves RA, and the net loan position, LA-LL. Absent valuation changes, unrequited capital transfers, debt forgiveness and other debt reduction operations, and ignoring mis-invoicing of current account transactions, the rate of change of NFA would just equal the current account balance CA:

$$\Delta NFA_{jt} = CA_{jt} \tag{A.2}$$

Given some initial condition for NFA, we can use equation (A.2) recursively to construct the country's net foreign asset position. Analogously, we can construct each of the stocks in (A.1) by accumulating the disaggregated financial account flows. However, the conditions under which historical flow accumulation would yield a good approximation to the value of the corresponding stocks are quite stringent. Specifically, problems with valuation effects might arise from: (i) cross-exchange rates, whose effect depends on the currency composition of foreign assets and liabilities (generally unavailable from standard sources), and, (ii) changes in the secondary-market price of assets (e.g. equity prices in the case of portfolio investment, or market prices of developing country debt).[30]

The valuation problems would be overcome if there were available information on asset stocks at current exchange rates and market prices. However, this information is limited to two main sources: (i) the foreign reserve data collected by the IMF's International Financial Statistics (IFS), which value foreign exchange reserves at current exchange rates and have very broad coverage across countries over time; and (b) the external debt data compiled by the World Bank and OECD for most developing countries starting in 1970, which report debt at face value (after adjusting for debt forgiveness and reduction as well as changes in exchange rates). In addition, we have also the international investment positions (IIPs) of the IMF's Balance of Payments, which cover the majority of industrial countries over a varying number of years since the 1980s as well as a

[30]The latter valuation effects are even more difficult to estimate, as organized secondary markets often do not exist (particularly in developing countries).

handful of developing economies. The valuation methods underlying the BOP's IIPs vary across countries, as well as over time for a given country.

Given these facts, we take as primary data sources the IMF's IFS and BoP and the World Bank's Global Development Finance, complemented in a few cases by country-specific documents, typically from the respective central banks, plus the data on international investment positions constructed by Rider (1994) primarily for industrial countries. From these sources, we construct our foreign asset and liability stocks as follows (see Kraay et al. 2000 for more details). For reserves of all countries, as well as developing country debt liabilities, we simply take the values reported by the IMF and the World Bank, respectively. For all other assets and liabilities, we construct stock series from the flows reported by the BoP, using the earliest available stock (if one exists) to tie down the level of the series.[31] From these initial values, stock series are obtained using the recursive formula:

$$S_t = \frac{Q_t}{Q_{t-1}} S_t (1 - d) + F_t \tag{A.3}$$

where S denotes the dollar value of the stock at the end of the period, F is the net flow during the period, Q is the market price of the asset in current US dollars, and δ is the rate of physical depreciation. The key issue concerns the measurement of Q. In the case of FDI, we take Q to follow the replacement value of physical capital. For inward FDI, this is captured by the investment deflator of the host country. For outward FDI, a detailed breakdown of flows by destination is not available, and hence we use a weighted average of investment deflators, with weights given by the structure of intra-OECD flows; we set d at 4 percent. For portfolio equity liabilities, we set $d=0$ and measure Q by the domestic stock market price index (in US dollars), when one is available; otherwise, we use the same valuation as for FDI liabilities. In turn, for portfolio assets -whose breakdown across debtors is unavailable- we take Q to equal the Morgan-Stanley world stock market index.

So far we have ignored the problem of mis-measurement of capital flows and stocks. To attempt to capture unrecorded (net) assets, we augment our measure of recorded non-equity assets LA by adding to it the cumulative errors and omissions of the Balance of Payments, starting from the earliest data for which the information is available. By the very nature of unrecorded assets, it is impossible to know their composition by currency and type of financial instrument, so that in this case we do not attempt to introduce any valuation adjustment.

[31] For most countries, initial FDI stocks are obtained from OECD (1967), which reports direct investment assets of each industrial country disaggregated by country of destination; this provides also the basic source of initial values for developing country inward FDI. For portfolio equity assets and liabilities, stock information is generally not available, although this is not too serious a problem given that portfolio flows are a relatively recent phenomenon; absent an initial stock, we set the starting value at zero. For industrial country loan assets and liabilities, as well as for developing country loan assets, we take as initial stocks those reported by the BoP, Rider's (1994) data, or national sources whenever available.

Appendix B: Estimating Panel Cointegration Models[32]

B.1. Model and Assumptions

Consider an *m*-vector Y_{it} of I(1) processes and an *m*-vector u_t of stationary time series whose long run covariance matrix (given by the value of the spectral density of u_{it} at zero) is non-singular. We partition these vectors as follows:

$$Y_{it} = \begin{bmatrix} y_{it} \\ x_{it} \end{bmatrix}_{p}^{1}; \quad u_{it} = \begin{bmatrix} u_{1it} \\ u_{2it} \end{bmatrix}_{p}^{1}; \quad m = 1 + p$$

and we assume that $\{Y_{it}\}$ is generated by the cointegrated system:

$$\begin{aligned} y_{it} &= \mu_i + x_{it}'\beta + u_{1it} \\ x_{it} &= x_{i,t-1} + u_{2it} \end{aligned} \qquad \text{(B.1)}$$

where β is the *px1* vector of slope parameters and the first equation in (B.1) is interpreted as the stochastic version of a linear long-run equilibrium relationship and u_{it} represents the stationary deviations from equilibrium. This equation corresponds to a fixed-effects panel regression, with μ_i being the intercepts. On the other hand, the second equation in (B.1) specifies x_{it} as a *px1* integrated processes of order one, for all *i*.

Assumption 1. We use the sequential limit theory (*i.e.* $T\rightarrow\infty$ is followed by $N\rightarrow\infty$) developed by Phillips and Moon (1999) as the basis of our asymptotic theory.[33]

Assumption 2. The innovation vector $u_t = (u_{1it}, u_{2it}')'$, is a linear process that satisfies, for each *i*, the following: (a) $u_t = C(L)\varepsilon_{it} = \sum_{j=0}^{\infty} c_j\varepsilon_{it-j}$, where $\sum_{j=0}^{\infty} j^a \left|C_j\right| < \infty$, and $\left|C(1)\right| \neq 0$ for some a>1; (b) ε_{it} is *iid* with zero mean, variance matrix $\sum_{\varepsilon}$, and finite fourth order cumulants.

In order to obtain the relevant asymptotics, we use the fact that the innovation process u_t satisfies the invariance principle (IP).[34] Hence, from assumption 2, the partial sum process $\frac{1}{\sqrt{T}}\sum_{t=1}^{[Tr]} u_{it}$ satisfies the following multivariate invariance principle (Phillips and Solo, 1992):

$$\frac{1}{\sqrt{T}}\sum_{t=1}^{[Tr]} u_{it} \Rightarrow B_i(r) = \begin{bmatrix} B_{1i} \\ B_{2i} \end{bmatrix} \equiv BM_i(\Omega), \quad T \rightarrow \infty, \forall i \qquad \text{(B.2)}$$

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[32]The appendix draws heavily from Phillips and Moon (1999), Kao (1999), and Kao and Chiang (1999).

[33]According to Kao (1999), one of the main issues in deriving the asymptotic distribution of the double indexed model is how to treat the two indices, *T* and *N*. In general, there are three types of convergence: (i) sequential limit, $T\rightarrow\infty$ first and then $N\rightarrow\infty$, (ii) pair-wise limit, assuming N(T) where $N(T) \rightarrow\infty$ as $T\rightarrow\infty$, and (iii) joint limit, N and $T\rightarrow\infty$. The joint limit is the strongest concept among the three.

[34]To achieve asymptotic normality using the *(N,T)*-asymptotics, we need to make strong assumptions. For example, we assume that the error terms are independent across countries.

The long run covariance matrix of u_{it} is:

$$\Omega = \sum_{j=-\infty}^{\infty} E\left[u_{it}u_{i0}'\right] = C(1)\Sigma_e C(1)' = \Sigma + \Gamma + \Gamma' = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix}$$

where:

$$\Gamma = \sum_{j=1}^{\infty} E\left[u_{it}u_{i0}'\right] = \begin{bmatrix} \Gamma_{11} & \Gamma_{12} \\ \Gamma_{21} & \Gamma_{22} \end{bmatrix}; \quad \Sigma = \sum_{j=1}^{\infty} E\left[u_{i0}u_{i0}'\right] = \begin{bmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{bmatrix}$$

are partitioned conformably with u_{it}.

Assumption 3. The matrix Ω_{22} is non-singular, *i.e.* x_{it} are not cointegrated.

We define: $\Omega_{1\cdot 2}=\Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21}$. Then, we can rewrite B_i as:

$$B_i = \begin{bmatrix} B_{1i} \\ B_{2i} \end{bmatrix} = \begin{bmatrix} \Omega_{1\cdot 2}^{\ -1} & \Omega_{12}\Omega_{22}^{-1} \\ 0 & \Omega_{22}^{1/2} \end{bmatrix} \begin{bmatrix} M_{1i} \\ M_{2i} \end{bmatrix}$$

where $(M_{1i}\ M_{2i}')' \equiv BM(I)$ is a standardized Brownian motion.

Now, we define the one-sided long run covariance:

$$\Delta = \sum_{j=0}^{\infty} E\left[u_{ij}u_{i0}'\right] = \Sigma + \Gamma = \begin{bmatrix} \Delta_{11} & \Delta_{12} \\ \Delta_{21} & \Delta_{22} \end{bmatrix}$$

Throughout this appendix we assume that our panels are homogeneous (*i.e.* variances are constant across the cross-section units). Finally, following assumption 1, all limits in the upcoming theorems are taken as T ®¥ followed by N ®¥ sequentially.

B.2. The Asymptotics of the OLS Estimator

If we apply OLS to the cointegrating relationship in the system (B.1), we find that:

$$\hat{b}_{OLS} = \left[\sum_{i=1}^{N}\ \sum_{t=1}^{T}(x_{it}-\bar{x}_i)(x_{it}-\bar{x}_i)'\right]^{-1}\left[\sum_{i=1}^{N}\ \sum_{t=1}^{T}(x_{it}-\bar{x}_i)(y_{it}-\bar{y}_i)'\right] \qquad \text{(B.3)}$$

Theorem 1 *(Kao and Chiang, 1999)*. If assumptions 1-3 hold, then:

(a) $T(\hat{b}_{OLS}-\beta) \rightarrow_p -3\Omega_{22}^{-1}\Omega_{21} + 6\Omega_{22}^{-1}\Delta_{21}$

(b) $\sqrt{N}\ T(\hat{b}_{OLS}-\beta) - \sqrt{N}\ \delta_{NT} \Rightarrow N(0, \Omega_{22}^{-1}\Omega_{1\cdot 2})$

where:

$$d_{NT} = \left[\frac{1}{N}\sum_{i=1}^{N}\frac{1}{T^2}\sum_{t=1}^{T}(x_{it}-\bar{x}_i)(x_{it}-\bar{x}_i)'\right]^{-1}\left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{it}-\bar{x}_i)(y_{it}-\bar{y}_i)'\right]$$

and $\tilde{M}_{2i} = M_{2i} - \int M_{2i}$.[35]

[35]The asymptotic covariance matrix in part (b) of the Theorem can be interpreted as long run noise-to-signal ratio $\Omega_{22}^{-1}\Omega_{1\cdot 2}$. In addition, the term $-(1/2)\Omega_{21}$ is attributed to the endogeneity of the regressor x_{it}, whereas Δ_{21} is due to serial correlation. Finally, we can show that $\delta_{NT} \rightarrow_p -3\Omega_{22}^{-1}\Omega_{21}+6\Omega_{22}^{-1}\Delta_{21}$.

We define $\hat{\Omega}_{22}$, $\hat{\Omega}_{21}$, and $\hat{\Delta}_{21}$ as consistent estimators of Ω_{22}, Ω_{21}, and Δ_{21}, respectively. Then, using part (b) in Theorem 1, we define the bias-corrected OLS estimator, $\hat{\boldsymbol{b}}^{+}_{OLS}$, as:

$$\hat{\boldsymbol{b}}^{+}_{OLS} = \hat{\boldsymbol{b}}_{OLS} - \frac{1}{T}\hat{\boldsymbol{d}}_{NT} \tag{B.4}$$

such that:

$$\sqrt{N}T(\hat{\boldsymbol{b}}^{+}_{OLS} - \boldsymbol{b}) \Rightarrow N(0,6\Omega_{22}^{-1}\Omega_{1\cdot 2}) \tag{B.5}$$

where $\hat{\boldsymbol{d}}_{\mathrm{NT}} = -3\hat{\Omega}_{22}{}^{-1}\hat{\Omega}_{21} + 6\hat{\Omega}_{22}{}^{-1}\hat{\Delta}_{21}$.

B.3. The Dynamic OLS Estimator (DOLS)

The dynamic OLS estimator, $\hat{\boldsymbol{b}}_{\mathrm{DOLS}}$, uses past and future values of Δx_{it} as additional regressors. Before we proceed to show the limiting distribution of the $\hat{\boldsymbol{b}}_{\mathrm{DOLS}}$, we need the following additional assumption.

Assumption 4. The spectral density matrix $f_{ww}(\lambda)\geq\delta I_T$, $\lambda\in[0,\pi]$, $\delta>0$.

When assumptions 2 and 4 hold, the process u_{1it} can be written as (Saikkonen, 1991):

$$u_{1it} = \sum_{j=-\infty}^{\infty} \boldsymbol{J}_{ij} u_{2i,t+j} + v_{it}; \quad \forall i \tag{B.6}$$

where $\sum_{j=-\infty}^{\infty}\left\|\boldsymbol{J}_{ij}\right\|<\infty$, *i.e.* $\{\vartheta_{ij}\}$ are assumed to be absolutely summable. The process $\{v_{it}\}$ is stationary with zero mean, $\{v_{it}\}$ and $\{u_{2it}\}$ are uncorrelated not only contemporaneously but also in lags and leads. In practice, the leads and lags may be truncated while retaining (B.6) approximately, so that

$$u_{1it} = \sum_{j=-q}^{q} \boldsymbol{J}_{ij} u_{2i,t+j} + v_{it}; \quad \forall i \tag{B.7}$$

In addition, we require that q tends to infinity with T at a suitable rate.

Assumption 5. q ®¥ as T ®¥ such that q^3/T ® 0, and $\sqrt{T}\sum_{|j|>q}\left\|\boldsymbol{J}_{ij}\right\| \to 0, \forall i$.

We substitute (B.7) into first equation of the system specified in (B.1) to get:

$$y_{it} = \boldsymbol{m}_i + x_{it}'\boldsymbol{b} + \sum_{j=-q}^{q} \boldsymbol{J}_{ij} u_{2i,t+j} + \dot{v}_{it} \tag{B.8}$$

where $\dot{v}_{it} = v_{it} + \sum_{|j|>q} \boldsymbol{J}_{ij} u_{2i,t+j}$. Therefore, we obtain the dynamic OLS estimator ($\hat{\boldsymbol{b}}_{\mathrm{DOLS}}$) by running the following regression:

$$y_{it} = \boldsymbol{m}_i + x_{it}'\boldsymbol{b} + \sum_{j=-q}^{q} \boldsymbol{J}_{ij}\Delta x_{i,t+j} + \dot{v}_{it} \tag{B.9}$$

Theorem 2 *(Kao and Chiang, 1999)*. If assumptions 1-5 hold, then:

$$\sqrt{N}T(\hat{\boldsymbol{b}}_{DOLS} - \boldsymbol{b}) \Rightarrow N(0, 6\Omega_{22}^{-1}\Omega_{1\cdot 2})$$

Proof. First, we write (B.9) in vector form:

$$y_i = e\boldsymbol{a}_i + x_i\boldsymbol{b} + Z_{iq}C + \dot{v}_i \;^{o}\; x_i\boldsymbol{b} + Z_iD + \dot{v}_i$$

where y_i is a *Tx1* vector of y_{it}, e is a *Tx1* unit vector, Z_{iq} is a *Tx2q* matrix of observations on the *2q* regressors, $\Delta x_{i,t-q}$, ..., $\Delta x_{i,t+q}$; x_i is a *Txk* matrix of x_{it}, C is a *(2q)x1* vector of ϑ_{ij}, $\dot{v}_i$ is a *Tx1* vector of $\dot{v}_{it}$, Z_i is a *Tx(2q+1)* is the matrix $Z_i = (e, Z_{iq})$, and D is the *(2q+1)x1* vector of parameters.

Let $Q_i = I - Z_i(Z_i'Z_i\}^{-1}Z_i'$. It follows that:

$$(\hat{\boldsymbol{b}}_{DOLS} - \boldsymbol{b}) = \left[\sum_{i=1}^{N}(x_i'Q_ix_i)\right]^{-1}\left[\sum_{i=1}^{N}(x_i'Q_i\dot{v}_i)\right]$$

We rescale ($\hat{\boldsymbol{b}}_{DOLS}$ -β) by $\sqrt{N}$ T to get:

$$\sqrt{N}T(\hat{\boldsymbol{b}}_{DOLS} - \boldsymbol{b}) = \left[\frac{1}{N}\sum_{i=1}^{N}\frac{1}{T^2}(x_i'Q_ix_i)\right]^{-1}\left[\sqrt{N}\frac{1}{N}\sum_{i=1}^{N}\frac{1}{T}(x_i'Q_i\dot{v}_i)\right]$$

Following Saikkonen (1991), we can show that (as T→∞, for all *i*):

$$\frac{1}{T^2}(x_i'Q_ix_i) = \frac{1}{T^2}(x_i'W_Tx_i) + o_p(1) = \frac{1}{T^2}\sum_{t=q+1}^{T-q}(x_{it} - \bar{x}_i)(x_{it} - \bar{x}_i)' + o_p(1) \Rightarrow \int\tilde{B}_{2i}\tilde{B}_{2i}'$$

and

$$\frac{1}{T}(x_i'Q_i\dot{v}_i) = \frac{1}{T}(x_i'W_T\dot{v}_i) + o_p(1) = \frac{1}{T}\sum_{t=q+1}^{T-q}(x_{it} - \bar{x}_i)\dot{v}_{it} + o_p(1) \Rightarrow \int\tilde{B}_{2i}\,dB_{1\cdot 2}$$

where $\tilde{B}_{2i} = B_{2i} - \int B_{2i}$, $W_T = I_T - (1/T)ee'$, and $B_1 = \Omega_{12}\Omega_{22}^{-1}B_2 + B_{1\cdot 2}$.

Following Kao (1999), we can show that $\frac{1}{\sqrt{N}}\sum_{i=1}^{N}\int\tilde{B}_{2i}\tilde{B}_{2i}' = E\left[\int\tilde{B}_{2i}\tilde{B}_{2i}'\right] \to \frac{1}{6}\Omega_{22}$. On the other hand, if we apply the multivariate Lindeberg-Levy Central Limit Theorem, we have $\frac{1}{\sqrt{N}}\sum_{i=1}^{N}\int\tilde{B}_{2i}\,dB_{1\cdot 2i} \Rightarrow N(0, \frac{1}{6}\Omega_{22}^{-1}\Omega_{1\cdot 2})$ and, hence:

$$\sqrt{N}T(\hat{\boldsymbol{b}}_{DOLS} - \boldsymbol{b}) \Rightarrow N(0, 6\Omega_{22}^{-1}\Omega_{1\cdot 2})$$

Appendix C: Testing for Structural Change[36]

Consider the system of cointegrated regressions, as specified in (B.1):

$$y_{it} = \mu_i + x_{it}'\beta_t + u_{1it} \qquad \text{(C.1)}$$
$$x_{it} = x_{i,t-1} + u_{2it}$$

where β_t is a *px1* vector of slope parameters. As specified in (B.2), we assume that the innovation process u_t have zero means and satisfy the invariance principle. In addition, we assume that the $\{u_{it}\}$ processes are independent across *i*, and that $\{x_{it}\}$ are not cointegrated (*i.e.* Ω_{22} is non-singular, as in assumption 3 in Appendix B). Finally, we assume homogeneous panels (*i.e.* constant variances across countries).

The main goal is to test changes in the parameter vector β_t, where the change points are unknown. Hence, we consider the alternative hypothesis that there is only one change point (t_B) over time,

$$H_1 : \boldsymbol{b}_t = \begin{cases} \boldsymbol{b}_1; \; for\ t = 1,...,t_B \\ \boldsymbol{b}_2; \; for\ t = t_{B+1},...,T \end{cases} \qquad \text{(C.2)}$$

Under (C.2), we can rewrite (C.1) as:

$$\begin{aligned} y_{it} &= \mu_i + x_{it}'I(t \le t_B)\beta_1 + x_{it}'I(t > t_B)\beta_2 + u_{1it} \qquad \text{(C.3)} \\ &= \mu_i + x_{it}'(t_B)\beta + u_{1it} \end{aligned}$$

where $\beta = (\beta_1', \beta_2')'$, $I(\cdot)$ is an indicator function, and

$$x_{it}(t_B) = \begin{cases} x_{it} \cdot I(t \le t_B) \\ x_{it} \cdot I(t > t_B) \end{cases}$$

We define the average of the explanatory variables as:

$$\bar{x}_i(t_B) = \begin{cases} \dfrac{1}{T}\displaystyle\sum_{t=1}^{T} x_{it} \cdot I(t \le t_B) \\ \dfrac{1}{T}\displaystyle\sum_{t=1}^{T} x_{it} \cdot I(t > t_B) \end{cases}$$

The fixed effect representation of (C.3) is:

$$y_{it}^* = x_{it}^*(t_B)'\beta + u_{1it}^* \qquad \text{(C.4)}$$

where $y_{it}^* = y_{it} - \bar{y}_i$, $x_{it}^*(t_B) = x_{it}(t_B) - \bar{x}_i(t_B)$, and $u_{1it}^* = u_{1it} - \bar{u}_{1i}$. Now we define, $\hat{u}_{1it}^*(t_B) = y_{it}^* - x_{it}^*(t_B)'\hat{\boldsymbol{b}}^*(t_B)$ and $S(t_B) = \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{u}_{1it}^*(t_B)^2$, where $\hat{\boldsymbol{b}}^*(t_B)$ can be estimated using the dynamic least squares technique presented in Appendix B. On the other hand, the least squares estimate of t_B is defined as:

[36] The tests developed in this section are taken from Kao and Chiang (2000).

$$\hat{t}_B = \underset{t_B}{\arg\min}\; S(t_B) \tag{C.5}$$

According to Theorem 2 (Kao and Chiang, 1999) presented in Appendix B, $\sqrt{N}T(\hat{\boldsymbol{b}}_{DOLS} - \boldsymbol{b}) \Rightarrow N(0, 6\Omega_{22}^{-1}\Omega_{1\cdot 2})$ and we would like to test parameter stability in equation (C.1) with the alternative hypothesis (C.2). That is, we want to test $H_0: \beta_1 = \beta_2$ *vs.* $H_1: \beta_1 \neq \beta_2$.

Kao and Chiang (2000) propose a Wald statistic to test the null of parameter stability over time:

$$W(t_B) = \frac{1}{\hat{\Omega}_{1\cdot 2}}\left(\boldsymbol{b}_{1,t_B} - \boldsymbol{b}_{2,t_B}\right)'\left[(X_1'X_1)^{-1} + (X_2'X_2)^{-1}\right]^{-1}\left(\boldsymbol{b}_{1,t_B} - \boldsymbol{b}_{2,t_B}\right) \tag{C.6}$$

where $X_1 = X\cdot I(t \leq t_B) - \bar{X}\cdot I(t \leq t_B)$, $X_2 = X\cdot I(t > t_B) - \bar{X}\cdot I(t > t_B)$, and $\hat{\Omega}_{1\cdot 2}$ is a consistent estimate of $\Omega_{1\cdot 2}$ under the null (Kao and Chiang, 1999).

Assumption 6. We assume that $t_B/T \to r$, and $t_B, T \to \infty$.

Theorem 3 *(Kao and Chiang, 2000)*. If all the assumptions stated above and H_0 hold, then:

$$W(t_B) \Rightarrow Q_p(r) \equiv \frac{1}{r^2 + (1-r)^2}\left[B((1-r)^2) - B(r^2)\right]'\left[B((1-r)^2) - B(r^2)\right]$$

uniformly in r.

Remark *(Kao and Chiang, 2000)*. [1] The limiting distribution of W(k) is nuisance-parameter-free and depends only on the number of regressors, p.

[2] Since $B((1-r)^2)-B(r^2)$ has variance $(1-r)^2 - r^2$, $B((1-r)^2)-B(r^2)$ is a Bessel process of order p. Then, we can rewrite

$$Q_p(r) = \frac{(1-r)^2 - r^2}{(1-r)^2 + r^2}\frac{\left[B((1-r)^2) - B(r^2)\right]'\left[B((1-r)^2) - B(r^2)\right]}{(1-r)^2 - r^2} = \frac{(1-r)^2 - r^2}{(1-r)^2 + r^2}\frac{BM(s)'BM(s)}{s}$$

where the second term of the multiplication on the left-hand side is the square of a standardized Bessel process, $s=(1-r)^2 - r^2$, and $BM(s)$ denotes a p-vector of independent Brownian processes on $[0,\infty]$. Note that $r \neq 1/2$ since $s=0$ if $r=1/2$.

[3] The limiting distribution of $\frac{(1-r)^2 + r^2}{(1-r)^2 - r^2}W(t_B)$ has the same form as Andrews (1993).

Consider now the following statistics:

$$\sup W(t_B) = \sup_{[Tr^*] \leq t_B \leq T - [Tr^*]} W(t_B)$$

$$meanW(t_B) = \frac{1}{T}\sum_{t_B=[Tr^*]}^{T-[Tr^*]} W(t_B)$$

$$\exp W(t_B) = \ln\left(\frac{1}{T}\sum_{t_B=[Tr^*]}^{T-[Tr^*]} \exp\left(\frac{1}{2}W(t_B)\right)\right)$$

Using the continuous mapping theorem we then have the following corollary:

Corollary *(Kao and Chiang, 2000)*. Suppose the assumptions in Theorem 3 hold and under H_0:

$$\sup W(t_B) = \sup_{r^* \le r \le 1-r^*} Q_p(r)$$

$$meanW(t_B) = \int_{r^*}^{1-r^*} Q_p(r)dr$$

$$\exp W(t_B) = \ln\left(\int_{r^*}^{1-r^*} \exp\left(\frac{1}{2}Q_p(r)\right)dr\right)$$

The critical values for the test statistics sup $W(t_B)$, Mean $W(t_B)$, and exp$W(t_B)$ are provided by Kao and Chiang (2000).

Real Exchange Rates in the Long and Short Run: A Panel Co-Integration Approach

Appendix of Results

Table 1
Panel Unit Roots (Im, Pesaran and Shin, 1995): The $\bar{t}_{NT}$ Statistic

Variable	Period	Levels	Differences
Real Effective Exchange Rate, ln(q)	1960-97	-2.2005	-2.8350**
Ratio of Net Foreign Assets to GDP, (F/Y)	1966-97	-2.0634	-2.0177**
Traded Sector Productivity, $\ln(y_T/y_T^*)$	1960-97	-2.0069	-2.3067**
TFP Traded Sector, $\ln(A_T/A_T^*)$	1960-97	-2.1058	-2.5349**
Terms of Trade, $\ln(P_X/P_M)$	1960-97	-2.0147	-2.6887**
Non-Traded Productivity, $\ln(A_N/A_N^*)$	1965-97	-2.1438	-2.2291**

*Notes: Before performing the ADF regressions for individual countries, we remove the common time dummies from all variables. The ADF regression in levels includes the time trend, whereas the ADF regression in differences does not. In the latter case, the alternative hypothesis is that series is stationary around a constant since any time trend in levels will be removed by differencing. This table reports the t-bar ($\bar{t}_{NT}$) statistic, defined as the sample average of the t-statistics obtained from the ADF regressions of individual countries. For 67 countries during the 1960-97 period (32 time series observations), the approximate sample critical values of the $\bar{t}_{NT}$ statistic are: (i) Without deterministic trend: –1.78, -1.71, and –1.66 at the 1, 5, and 10 percent significance level; (ii) With deterministic trend: –2.41, -2.34, and –2.30 at the 1, 5, and 10 percent significance level. For more details, see Table 4 in Im, Pesaran and Shin (1995). * (**) indicates that the test is significant at the 10 (5) percent level.*

Table 2
The Cointegration Tests
Cointegration Relationship: The Equilibrium Real Exchange Rate Equation, 1966-97
Using different proxies for the productivity and time-periods

Cointegration Test	1966-97 Using y_T/y_T^*	1966-97 Using A_T/A_T^*	1973-97 Using y_T/y_T^*	1973-97 Using A_T/A_T^*
The Kao (1999) Panel Cointegration Tests				
DF_ρ	-9.0544**	-8.1962**	-8.5138**	-7.845**
DF_t	78.384**	78.9037**	78.6204**	79.0826**
DF_ρ^*	-14.9704**	-13.6543**	-13.6607**	-12.6455**
DF_t^*	-4.8236**	-4.4879**	-4.6986**	-4.3967**
ADF	-5.7358**	-5.512**	-5.594**	-5.1824**
The Pedroni (1995) Panel Cointegration Tests				
PC_1	-29.3103**	-27.269**	-28.5985**	-27.2707**
PC_2	-28.8487**	-26.8396**	-28.0207**	-26.7198**
No. Countries	67	67	67	67
No. Observations	2144	2144	1675	1675

Notes: The dependent variable is the real effective exchange rate and the explanatory variables are the ratio of net foreign assets to GDP, the ratio of Home to Foreign output per worker, the terms of trade, and the ratio of Home to Foreign non-traded output per worker. All variables are expressed in logs. The critical probabilities (p-values) are reported in parentheses. The cointegration test statistics are calculated through the residuals from the OLS estimation.

Table 3

Cross-Section Correlations between the Real Exchange rate and its determinants

Sample of Countries according to income and capital controls over the 1966-97 period

Variable	**All Countries**	**High Income**	**Low Income**	**Low Capital Controls**	**High Capital Controls**
I. Cross-Section Correlations, 1966-97					
Δ (F/Y)	0.0822	-0.0905	0.0408	0.1878	-0.0333
$\Delta\ln(y_T/y_T*)$	0.3280**	0.2785**	0.3281**	0.4256**	0.1699
$\Delta\ln(P_X/P_M)$	0.3202**	0.0462	0.4458**	0.4708**	0.1110
$\Delta\ln(A_N/A_N*)$	0.2602**	0.2316*	0.2826*	0.5318**	0.0017
Conditional Correlation (Controlling for Demand Shocks)					
$\Delta\ln(A_N/A_N*)$	-0.0288	-0.0346	0.0859	0.1705	-0.1636
II. Panel Data Correlations					
Δ (F/Y)	-0.0254	0.0604	-0.0645	0.0287	-0.0828
$\Delta\ln(y_T/y_T*)$	0.2483**	0.4182**	0.1767**	0.2546**	0.2386**
$\Delta\ln(P_X/P_M)$	0.2382**	0.2624**	0.2371**	0.1983**	0.2744**
$\Delta\ln(A_N/A_N*)$	0.1784**	0.3412**	0.1205	0.3125**	0.0472
Conditional Correlation (Controlling for Demand Shocks)					
$\Delta\ln(A_N/A_N*)$	-0.1209**	-0.0930	-0.1487**	0.1466*	-0.3757**
III. Cross-Section Correlations over Decades (All Countries)					
	1966-97	**1966-73**	**1974-81**	**1982-89**	**1990-97**
Δ (F/Y)	0.0822	-0.0659	-0.0121	-0.0023	0.1470
$\Delta\ln(y_T/y_T*)$	0.3280**	0.1140	0.0656	0.3417**	0.2735**
$\Delta\ln(P_X/P_M)$	0.3202**	0.1396	0.2305**	0.4781**	0.1137
$\Delta\ln(A_N/A_N*)$	0.2602**	0.2901**	-0.0864	0.2507**	-0.0711
Conditional Correlation (Controlling for Demand Shocks)					
$\Delta\ln(A_N/A_N*)$	-0.0288	0.2265*	-0.3255**	-0.1128	-0.3580**

The panel data correlations are computed over a sample of 8-year period observations in the sample of 67 countries: 1966-73, 1974-81, 1982-89, and 1990-97. The conditional correlations are partial correlations between real exchange rates and non-traded productivity conditional on the evolution of consumption in the countries across the world.

** (**) denotes that the correlation coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 4
Estimating the Cointegration Relationship: The Equilibrium Real Exchange Rate Equation, 1966-97 (Using different proxies for the productivity and time-periods)
Estimation Method: Dynamic OLS (DOLS)

Variables	1966-97 Using y_T/y_T*	1973-97 Using y_T/y_T*	1966-97 Using A_T/A_T*	1973-97 Using A_T/A_T*
I. Ordinary Least Squares (OLS)				
(F/Y)	0.0017 (0.0088)	0.0112 (0.0092)	0.0073 (0.0104)	0.0264 (0.0103)**
ln(y_T/y_T*)	0.3579 (0.0430)**	-0.2238 (0.0454)**	0.2969 (0.0563)**	-0.4394 (0.0515)**
ln(P_X/P_M)	0.1366 (0.0248)**	0.1723 (0.0251)**	0.1547 (0.0343)**	0.2423 (0.0336)**
ln(A_N/A_N*)	-0.092 (0.0199)**	0.05 (0.0170)**	-0.0751 (0.0238)**	0.0679 (0.0193)**
R Squared	0.2524	0.2326	0.2339	0.2638
II. Ordinary Least Squares with Bias Correction				
(F/Y)	0.0076 (0.0149)	0.0145 (0.0152)	0.0175 (0.0176)	0.034 (0.0175)
ln(y_T/y_T*)	0.4185 (0.0943)**	-0.0685 (0.0943)	0.3241 (0.1184)**	-0.2673 (0.1024)**
ln(P_X/P_M)	0.125 (0.0465)**	0.1506 (0.0472)**	0.1238 (0.0567)**	0.2095 (0.0555)**
ln(A_N/A_N*)	-0.129 (0.0362)**	0.0185 (0.0321)	-0.1018 (0.0420)**	0.0328 (0.0356)
R Squared	0.2524	0.2326	0.2339	0.2638
III. Fully-Modified Estimator (FM-OLS)				
(F/Y)	0.15 (0.0154)**	0.1073 (0.0158)**	0.2011 (0.0183)**	0.1544 (0.0182)**
ln(y_T/y_T*)	1.4328 (0.0973)**	1.9601 (0.0909)**	1.5229 (0.1233)**	2.1151 (0.1067)**
ln(P_X/P_M)	0.606 (0.0480)**	0.4751 (0.0487)**	0.666 (0.0591)**	0.4934 (0.0578)**
ln(A_N/A_N*)	-0.3071 (0.0374)**	-0.0769 (0.0332)**	-0.4018 (0.0438)**	-0.1655 (0.0371)**
R Squared	0.3828	0.4387	0.3813	0.3821
IV. Dynamic Least Squares (DOLS)				
(F/Y)	0.2127 (0.0176)**	0.1658 (0.0181)**	0.2243 (0.0220)**	0.1734 (0.0219)**
ln(y_T/y_T*)	1.3024 (0.1118)**	1.8472 (0.1044)**	1.4871 (0.1480)**	2.1559 (0.1281)**
ln(P_X/P_M)	0.7427 (0.0551)**	0.5722 (0.0559)**	0.6077 (0.0710)**	0.5009 (0.0693)**
ln(A_N/A_N*)	-0.1837 (0.0429)**	-0.017 (0.0380)	-0.3565 (0.0525)**	-0.1651 (0.0445)**
R Squared	0.4262	0.4617	0.4129	0.4551
No. Countries	67	67	67	67
No. Observations	2144	2144	1675	1675

*Notes: 1/ The dependent variable is the real effective exchange rate and the explanatory variables are the ratio of net foreign assets to GDP, the ratio of Home to Foreign output per worker, the terms of trade, and the ratio of Home to Foreign non-traded output per worker. 2/ The Dynamic Least Squares (DOLS) estimations are performed with 2 lags and 1 lead, DOLS(2,1). The numbers in parenthesis represent the standard error of the estimators. * (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 5
Estimating the Cointegration Relationship: The Equilibrium Real Exchange Rate Equation
Sub-sample of countries according to income classification

	1966-97		1973-97	
Variables	**FM-OLS**	**DOLS**	**FM-OLS**	**DOLS**
I. High and Upper Middle Income Countries (33)				
(F/Y)	0.1566 (0.0115)**	0.1804 (0.0132)**	0.1734 (0.0115)**	0.1342 (0.0138)**
$\ln(y_T/y_T^*)$	3.9722 (0.1080)**	3.9263 (0.1240)**	4.1868 (0.1086)**	4.4664 (0.1304)**
$\ln(P_X/P_M)$	1.0534 (0.0630)**	1.3649 (0.0724)**	1.8366 (0.0586)**	1.8572 (0.0703)**
$\ln(A_N/A_N^*)$	-1.432 (0.0441)**	-1.4078 (0.0506)**	-1.6291 (0.0396)**	-1.8058 (0.0475)**
R Squared	0.4285	0.4318	0.4175	0.4029
No. Observations	1056	1056	825	825
II. Low and Lower Middle Income Countries (34)				
(F/Y)	0.2577 (0.0548)**	0.364 (0.0630)**	0.3132 (0.0654)**	0.3497 (0.0785)**
$\ln(y_T/y_T^*)$	0.8385 (0.1390)**	0.6953 (0.1597)**	0.9746 (0.1837)**	0.886 (0.2205)**
$\ln(P_X/P_M)$	0.1408 (0.0607)**	0.0045 (0.0698)	-0.1911 (0.0812)**	-0.2657 (0.0975)**
$\ln(A_N/A_N^*)$	0.2381 (0.0498)**	0.389 (0.0572)**	0.0831 (0.0630)	0.1788 (0.0755)**
R Squared	0.3904	0.4115	0.3799	0.3876
No. Observations	1088	1088	850	850
III. Testing the Equality of Coefficients between High and Low Income Countries				
	(F/Y)	$\ln(y_T/y_T^*)$	$\ln(P_X/P_M)$	$\ln(A_N/A_N^*)$
1966-97	1.4643 (0.2262)	42.1130 (0.0000)	12.6001 (0.0003)	35.9385 (0.0000)
	Overall Test:		61.5235 (0.0000)	
1973-97	1.3297 (0.2489)	39.4564 (0.0000)	10.9526 (0.0009)	51.8934 (0.0000)
	Overall Test:		53.2674 (0.0000)	

*Notes: 1/ 2/ See corresponding footnotes in Table 4. 3/ The income classification of countries follows the criteria proposed by the World Bank. * (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 6
Estimating the Cointegration Relationship: The Equilibrium Real Exchange Rate Equation
Sub-sample of countries according to presence of capital controls

Variables	1966-97 FM-OLS	1966-97 DOLS	1973-97 FM-OLS	1973-97 DOLS
I. Sample of Countries with Low Capital Controls (37)				
(F/Y)	0.1308 (0.0127)**	0.1635 (0.0146)**	0.182 (0.0161)**	0.1705 (0.0194)**
$\ln(y_T/y_T^*)$	2.5733 (0.1012)**	2.6581 (0.1162)**	2.7232 (0.1335)**	2.8127 (0.1602)**
$\ln(P_X/P_M)$	0.1934 (0.0584)**	0.0454 (0.0671)	0.0315 (0.0735)	-0.2149 (0.0881)**
$\ln(A_N/A_N^*)$	-1.0256 (0.0403)	-1.0766 (0.0463)**	-1.1716 (0.0487)**	-1.2138 (0.0584)**
R Squared	0.4028	0.4746	0.3824	0.3956
No. Observations	1184	1184	925	925
II. Sample of Countries with High Capital Controls (30)				
(F/Y)	-0.4163 (0.0561)**	-0.3795 (0.0645)**	-0.3425 (0.0686)**	-0.4052 (0.0824)**
$\ln(y_T/y_T^*)$	0.698 (0.1506)**	0.5347 (0.1729)**	0.6988 (0.1934)**	0.581 (0.2321)**
$\ln(P_X/P_M)$	0.6361 (0.0652)**	0.8949 (0.0749)**	0.9006 (0.0796)**	0.7428 (0.0955)**
$\ln(A_N/A_N^*)$	-0.3416 (0.0567)**	-0.1719 (0.0651)**	-0.3442 (0.0668)**	-0.2299 (0.0802)**
R Squared	0.3065	0.3859	0.2621	0.3492
No. Observations	960	960	750	750
III. Testing Equality of Coefficients between Countries with High and Low Capital Controls				
	(F/Y)	$\ln(y_T/y_T^*)$	$\ln(P_X/P_M)$	$\ln(A_N/A_N^*)$
1966-97	0.5183 (0.4715)	62.5239 (0.0000)	3.1784 (0.0746)	25.7660 (0.0000)
	Overall Test:		139.2946 (0.0000)	
1973-97	0.4798 (0.4885)	51.9346 (0.0000)	2.8173 (0.0932)	19.6497 (0.0000)
	Overall Test:		100.1649 (0.0000)	

Notes: 1/ 2/ See corresponding footnotes in Table 4. 3/ Our proxy for capital controls are the dummy variables constructed by Grilli and Milesi-Ferreti (1995). These dummy variables take the value of 1 when the control is present and 0 otherwise. They capture multiple exchange rate practices, controls on current account transactions, control on capital account transactions, and surrogate export proceeds. Our measure of capital controls is the sum of these four dummies (i.e. it takes values from 0 to 4). If the average of this measure over the sample period of estimation is greater than 3, then we consider that the countries has high capital controls. Otherwise, the country has low capital controls.
** (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 7
Estimating the Cointegration Relationship: The Equilibrium Real Exchange Rate Equation
Sub-sample of countries according to the intensity of capital controls

	1966-97		**1973-97**	
Variables	**FM-OLS**	**DOLS**	**FM-OLS**	**DOLS**
I. Sample of Countries with Low Black Market Premium (46)				
(F/Y)	0.1326 (0.0097)**	0.1833 (0.0112)**	0.1844 (0.0111)**	0.1888 (0.0133)**
$\log(y_T/y_T^*)$	2.4278 (0.0727)**	2.3375 (0.0835)**	2.4857 (0.0860)**	2.6263 (0.1032)**
$\log(P_X/P_M)$	-0.3525 (0.0455)**	-0.3338 (0.0523)**	-0.2537 (0.0499)**	-0.5896 (0.0599)**
$\log(A_N/A_N^*)$	-0.982 (0.0279)**	-0.8649 (0.0320)**	-1.0696 (0.0298)**	-1.1634 (0.0358)**
R Squared	0.3514	0.3958	0.3332	0.3613
No. Observations	1472	1472	1150	1150
II. Sample of Countries with High Black Market Premium (21)				
(F/Y)	0.0055 (0.1060)	0.0365 (0.1207)	0.1175 (0.1190)	0.2623 (0.1427)*
$\log(y_T/y_T^*)$	-0.6308 (0.2350)**	-0.9983 (0.2697)**	-0.3603 (0.3012)	-0.7176 (0.3614)*
$\log(P_X/P_M)$	1.4272 (0.0882)**	1.5269 (0.1013)**	1.5001 (0.1144)**	1.3207 (0.1373)**
$\log(A_N/A_N^*)$	0.8611 (0.0898)**	1.1883 (0.1031)**	0.623 (0.1130)**	0.8977 (0.1356)**
R Squared	0.3469	0.3641	0.3495	0.3530
No. Observations	672	672	525	525
III. Testing Equality of Coefficients between Countries with High and Low BMP				
	(F/Y)	$\ln(y_T/y_T^*)$	$\ln(P_X/P_M)$	$\ln(A_N/A_N^*)$
1966-97	0.3766 (0.5394)	125.245 (0.0000)	23.3087 (0.0000)	64.1156 (0.0000)
	Overall Test:		209.4625 (0.0000)	
1973-97	0.2537 (0.6149)	96.2148 (0.0000)	19.7065 (0.0000)	51.2348 (0.0000)
	Overall Test:		189.3714 (0.0000)	

Notes: 1/ 2/ See corresponding footnotes in Table 4. 3/ Given that the dummy variable approach does not capture the intensity of the capital controls, we use the black market premium as our proxy. If the black market premium averages over 20 percent in the sample period, then we consider that the country has high capital controls. If the average black market premium is below 20 percent, then the country has low capital controls. We also try with 10 percent as our benchmark for high/low capital controls, and the results were similar.
** (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 8
Testing the Robustness of the Long-Run Coefficients across Countries
Sample of ALL countries (67) for different sub-periods

Variables	All Countries, 1966-97			All Countries, 1973-97		
	Pooled DOLS	Average DOLS	Hausman Test	Pooled DOLS	Average DOLS	Hausman Test
(F/Y)	0.2127 (0.0176)**	1.6380 (2.0896)	3.4652 [0.0626]	0.1658 (0.0181)**	2.6079 (1.7774)	9.8876 [0.0017]
$\ln(y_T/y_T^*)$	1.3024 (0.1118)**	3.2714 (1.2326)**	4.5309 [0.0333]	1.8472 (0.1044)**	5.0460 (1.2572)**	10.4293 [0.0012]
$\ln(P_X/P_M)$	0.7427 (0.0551)**	1.8091 (1.3382)	7.6340 [0.0057]	0.5722 (0.0559)**	0.7387 (1.2502)	5.0177 [0.0251]
$\ln(A_N/A_N^*)$	-0.1837 (0.0429)**	-0.5105 (0.5996)	2.2955 [0.1297]	-0.017 (0.0380)	-0.8990 (0.5890)	4.2330 [0.0396]
Overall H-Test			9.0878 [0.0589]			8.5772 [0.0726]

Notes: 1/ Estimation Method: Dynamic Least Squares (DOLS). The estimation is performed with 2 lags and 1 lead, DOLS(2,1). The numbers in parenthesis represent the standard error of the estimators. 2/ The average DOLS represents the average of the country-by-country DOLS estimation. In order to test the homogeneity of our long-run coefficients we test the equality of the A_DOLS estimator with the pooled DOLS estimator. We construct individual and joint Hausman-type tests. Both tests have a Chi-squared distribution, with 1 degree of freedom for the individual test and 4 degrees of freedom for the joint test. In this column, the numbers in brackets represent the p-values.
** (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 9
Testing the Stability and Equality of the Long-Run Coefficients across Groups of Countries
Sample of countries according to income per capita: High and Upper-Middle Income vs. Low and Lower-Middle Income Countries

Variables	High and Upper-Middle Income Countries (33)			Low and Lower-Middle Income Countries (34)		
	Pooled DOLS	Average DOLS	Hausman Test	Pooled DOLS	Average DOLS	Hausman Test
(F/Y)	0.1804 (0.0132)**	4.5138 (3.1329)	1.9132 [0.1666]	0.364 (0.0630)**	-1.1533 (1.0770)	12.9778 [0.0003]
$\ln(y_T/y_T^*)$	3.9263 (0.1240)**	3.3125 (1.8090)*	0.1146 [0.7350]	0.6953 (0.1597)**	3.2315 (0.6732)**	13.4358 [0.0002]
$\ln(P_X/P_M)$	1.3649 (0.0724)**	3.8221 (2.0564)*	1.4260 [0.2324]	0.0045 (0.0698)	-0.1446 (0.6411)	10.0535 [0.0015]
$\ln(A_N/A_N^*)$	-1.4078 (0.0506)**	-0.7964 (0.9258)	0.4348 [0.5096]	0.389 (0.0572)**	-0.2339 (0.2830)	4.6413 [0.0312]
Overall H-Test			1.4050 [0.8433]			12.1881 [0.0160]

*Notes: 1/ 2/ See corresponding footnotes in Table 8. 3/ The income classification of countries follows the criteria proposed by the World Bank. * (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 10
Testing the Stability and Equality of the Long-Run Coefficients across Groups of Countries
Sample of countries according to the presence and intensity of Capita Controls

Variables	**Low Capital Controls (37)**			**High Capital Controls (30)**		
	Pooled DOLS	Average DOLS	Hausman Test	Pooled DOLS	Average DOLS	Hausman Test
(F/Y)	0.1635 (0.0146)**	-0.4941 (0.6581)	0.9980 [0.3178]	-0.3795 (0.0645)**	4.2676 (3.8552)	11.4526 [0.0007]
$\ln(y_T/y_T^*)$	2.6581 (0.1162)**	1.9929 (0.7632)**	0.7424 [0.3889]	0.5347 (0.1729)**	4.8483 (1.8116)**	5.6184 [0.0178]
$\ln(P_X/P_M)$	0.0454 (0.0671)	1.8431 (1.3928)	1.6622 [0.1973]	0.8949 (0.0749)**	1.7672 (1.2708)*	10.4695 [0.0012]
$\ln(A_N/A_N^*)$	-1.0766 (0.0463)**	-0.5185 (0.3106)*	3.1591 [0.0755]	-0.1719 (0.0651)**	-0.5006 (0.9560)	9.1176 [0.0025]
Overall H-Test			1.4381 [0.8375]			10.3396 [0.0351]
	Low Black Premium (46)			**High Black Market Premium (21)**		
	Pooled DOLS	Average DOLS	Hausman Test	Pooled DOLS	Average DOLS	Hausman Test
(F/Y)	0.1833 (0.0112)**	2.5033 (2.2508)	1.0625 [0.3027]	0.0365 (0.1207)	-0.2573 (1.7367)	10.0285 [0.0015]
$\ln(y_T/y_T^*)$	2.3375 (0.0835)**	3.5816 (1.2300)**	1.0184 [0.3129]	-0.9983 (0.2697)**	2.5919 (1.2384)**	8.0242 [0.0046]
$\ln(P_X/P_M)$	-0.3338 (0.0523)**	2.5910 (1.5751)*	3.4444 [0.0635]	1.5269 (0.1013)**	0.0964 (0.8192)	3.0035 [0.0831]
$\ln(A_N/A_N^*)$	-0.8649 (0.0320)**	-0.6444 (0.4638)	0.2249 [0.6353]	1.1883 (0.1031)**	-0.2171 (0.8970)	12.4229 [0.0004]
Overall H-Test			1.6728 [0.7956]			9.0414 [0.0601]

*Notes: 1/ 2/ See corresponding footnotes in Table 8. 3/ The classification of countries according to the presence and intensity of capital controls is in line with the footnotes in Tables 6 and 7. * (**) denotes that the coefficient is significantly different from zero at the 10 percent (5 percent) level.*

Table 11
Testing for Structural Change in a Cointegrated Regression in Panel Data
Alternative Hypothesis: There is only one change point in year t_B

Sample	**Time Break (t_B)**	**sup W(t_B)**	**mean W(t_B)**	**exp W(t_B)**
All Countries	1973	15.83**	5.34**	5.54**
Sample of Countries according to income levels				
High Income	1973	164.84**	32.00**	78.96**
Low Income	1976	41.54**	8.04**	17.38**
Sample of Countries according to capital controls				
High Controls	1973	125.51**	26.40**	59.85**
Low Controls	1985	27.81**	4.82**	10.85**

*Notes: The procedure to detect the break point in our cointegrating relationship and the test statistics are explained briefly in Appendix C. The critical values for these test statistics are taken from Kao and Chiang (2000). Note that the distribution of these test statistics depends only on the number of regressors (i.e. in our case, that number is equal to 4). * (**) represents statistical significance at the 10 (5) percent level.*

Table 12
Short-Run Dynamics of the Real Exchange Rate: Speed of Adjustment (f_i) and Half-Life of Equilibrium Deviations (h_i)
Sample Period: 1973-97

Region	Statistic	Model ECM_1		Model ECM_2	
		f_i	h_i	f_i	h_i
All Countries	Average	-0.1919	3.61	-0.2116	3.28
	Median	-0.1609	4.31	-0.1890	3.67
Classification of Countries according to Output per capita (The World Bank)					
High and Upper-Middle Income	Average	-0.2090	3.32	-0.2419	2.87
	Median	-0.1959	3.54	-0.2484	2.79
Low and Lower-Middle Income	Average	-0.1753	3.95	-0.1830	3.79
	Median	-0.1370	5.06	-0.1452	4.77
Classification according to the presence of Capital Controls (Grilli and Milesi-Ferreti, 1995)					
Low Controls	Average	-0.1823	3.80	-0.2174	3.19
	Median	-0.1834	3.78	-0.2233	3.10
High Controls	Average	-0.2036	3.40	-0.2046	3.39
	Median	-0.1516	4.57	-0.1554	4.46
Classification according to the intensity of Capital Controls (i.e. Black Market Premium)					
Low BMP	Average	-0.1984	3.49	-0.2211	3.13
	Median	-0.1926	3.60	-0.2416	2.87
High BMP	Average	-0.1777	3.90	-0.1911	3.63
	Median	-0.1317	5.26	-0.1410	4.92

Notes: 1/ In order to explore the short-run dynamics, we estimate the error correction models ECM1 and ECM2, as specified in equations (29) and (30). Note that the difference between these two models is the inclusion of relative money supplies. 2/ We report both the speed of adjustment (or convergence) to the equilibrium rate, which is denoted by the parameter f, and the half life of the deviations from the equilibrium, $h=ln(0.5)/ln(1+f)$.

Figure 1: Average Half-Life of Real Exchange Rate Deviations from the equilibrium, 1973-97

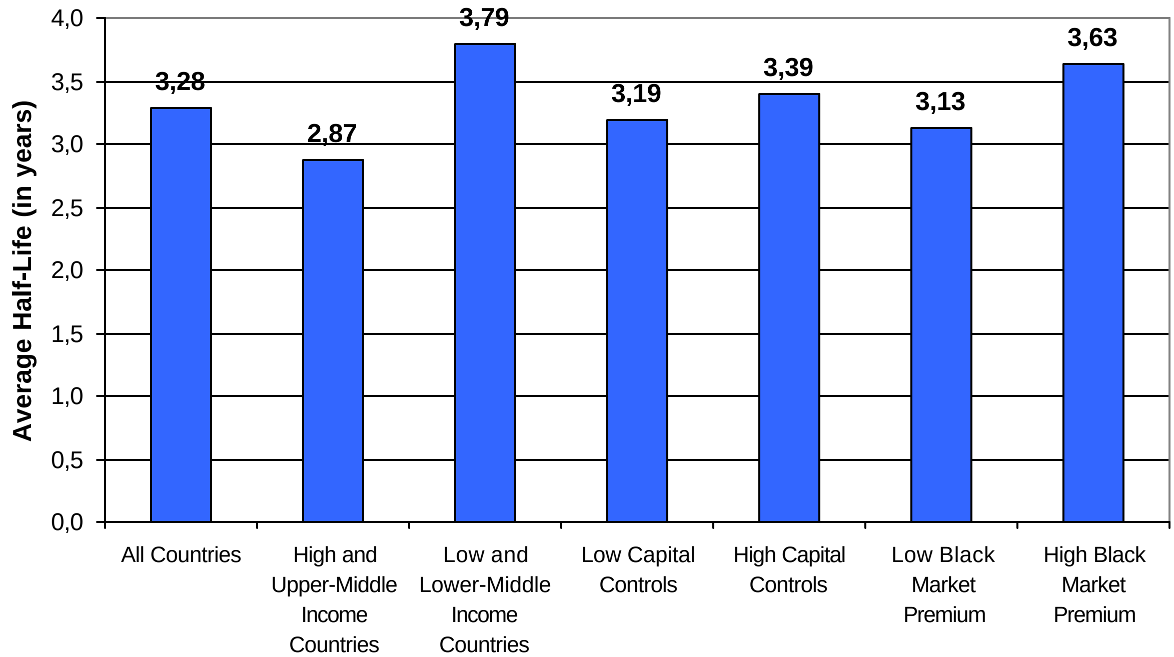


Figure 2: Median Half-Life of Real Exchange Rate Deviations from the equilibrium, 1973-97

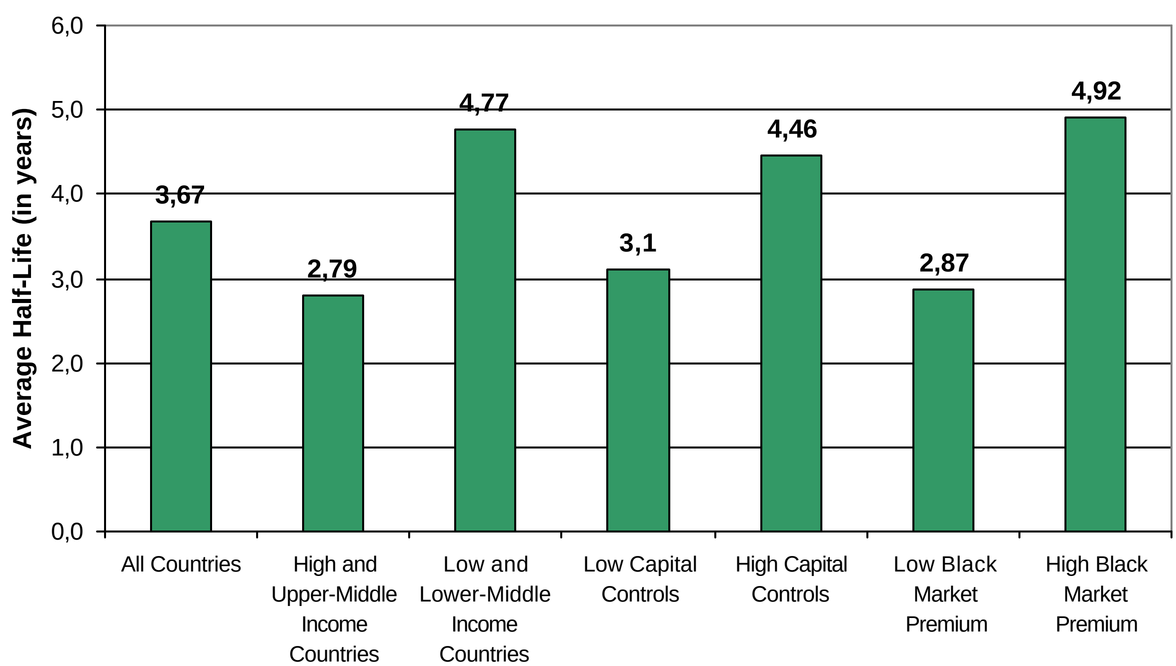


Figure 3: Conditional Correlation between the ratio of Income per Worker and the Half-Life of Real Exchange Rate Deviations

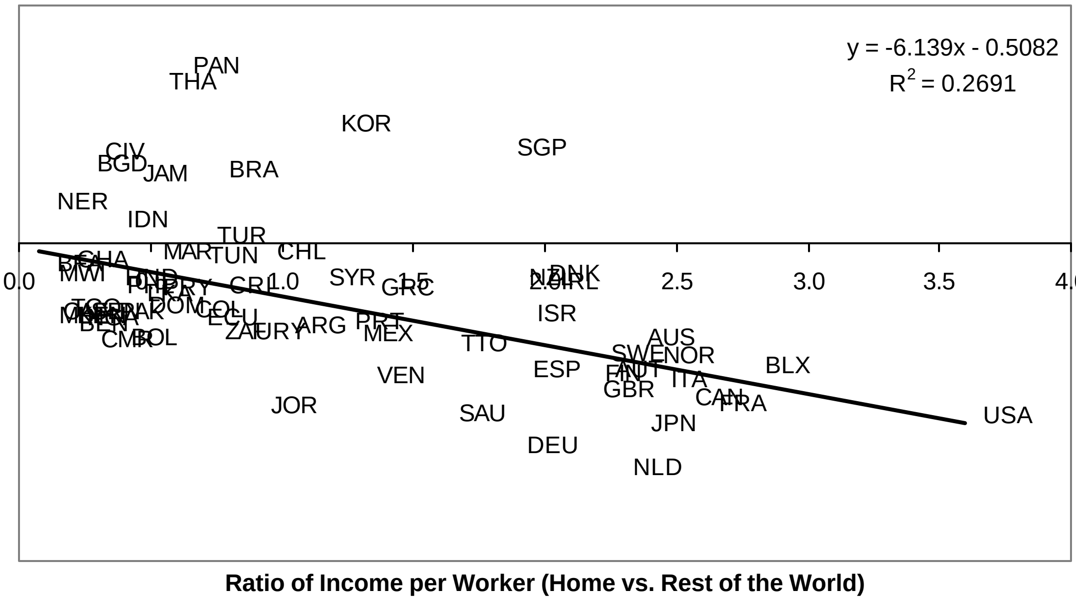
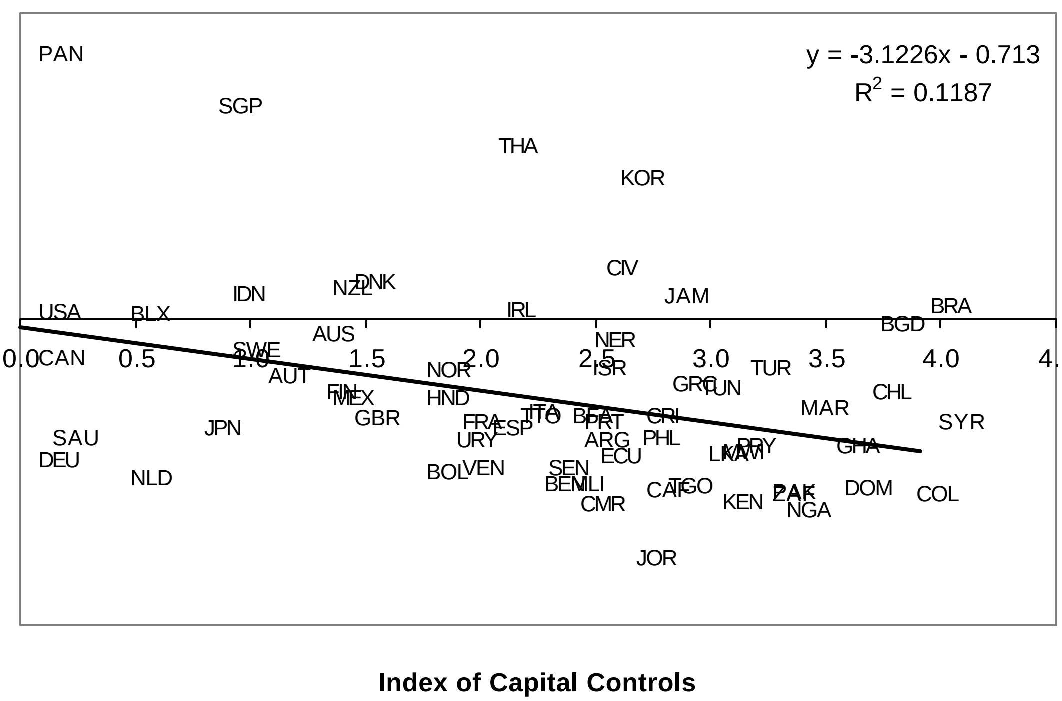


Figure 4: Conditional Correlation between the Index of Capital Controls and the Half-Life of Real Exchange Rate Deviations



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