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

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Abstract

The main goal of this paper is to tackle the empirical issues of the real exchange rate literature by applying recently developed panel cointegration techniques to a structural long-run real exchange rate equation. Using annual data for 67 countries over 1966-97, we find evidence of cointegration between the real exchange rate and its fundamentals. I also find: (a) evidence of cointegration holds for all sub-samples of countries (classified by income or capital controls), (b) parameter constancy across units holds only for high income countries and low capital controls, (c) structural change in the cointegrating relationship around 1973, (d) estimated parameters consistent with theoretical values implied with calibrated parameters of preferences and technology, (e) deviations from the equilibrium are large and persistent with half-life (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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I. Introduction

Recent advances in econometric testing and the availability of longer time series data have revitalized the PPP as a benchmark for evaluating long-run real exchange rates, although we still observe persistent PPP deviations (Froot and Rogoff, 1995; Wei and Parsley, 1995). The literature has usually linked these deviations 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. Empirically, the literature shows mixed evidence on the behavior of long-run real exchange rates. 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).

In order to tackle the empirical problems mentioned above, we estimate a long-run equilibrium relationship for the real exchange rate using the recently developed panel cointegration techniques (Kao, 1999; Kao and Chiang, 1999; Pedroni, 1999, Phillips and Moon, 1999). These 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, we formulate a simple real version of the traded vs. non-traded Obstfeld-Rogoff (1995) model in order to give a structural interpretation to our results.

We find evidence of a long-run relationship between real exchange rate and its fundamentals (net foreign assets, productivity in tradables and non-tradables, and the terms of trade), which is consistent with our model. This result holds for the full sample and all sub-samples of countries classified by income per capita and by capital controls. Robustness checks yield the following results: (1) long-run coefficient estimates are stable across countries for the samples of high-income countries and countries with low capital controls. (2) We find that the first oil crisis (1973) becomes a major source of instability over time in our long-run estimates for the sample of high-income countries and countries with low capital controls. (3) Our long-run estimates are consistent with calibrated theoretical values of our real exchange rate equation. However, the discrepancies are larger the higher is the value calibrated for the elasticity of intertemporal substitution. (4) Deviations from equilibrium real exchange rates decay at an average rate of *21.2* percent per year, with an implied half-life of *3.3* years for our full sample of countries, whereas the half-life of deviations for high-income countries is *2.87* years. These estimates lie near 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 II 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 III discusses the data used in our empirical evaluation. Section IV presents the panel data estimates for the long run real exchange rate. Section V introduces different robustness checks on our long-run estimates between and within countries as well as over time. Section VI discusses the behavior of real exchange rates in the short-run, thus computing half-lives of equilibrium real exchange rate deviations. Finally, section VII concludes.

II. The Theoretical Model

In order to give our empirical estimates a structural interpretation, we develop a simple model of real exchange rate behavior. Based on the traded vs. non-traded version of the Obstfeld and Rogoff (1995) model, we further assume the absence of money in the economy.[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, $\bar{y}_{T}$, 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 $[0, \mathrm{n}]$, whereas Foreign producers are located on $(n,1]$. We assume that all agents have similar preferences throughout the world over a real 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:

$$F_{t+1}^{j} = (1+r_t)F_t^{j} + p_{Nt}(j)y_{Nt}(j) - p_{Tt}^{X}\bar{y}_{Tt} - P_t C_t^{j} \tag{3}$$

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_{Tt}^{X} 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:

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

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

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:

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

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

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

Equation (7) reflects the consumption-based real interest-rate effect. That is, there is a shift towards present consumption as consumption-based real interest rate is lower if the aggregate price level relative to the price of tradables is currently low relative to its future value. However, it also encourages substitution from traded to non-traded goods. The inter-temporal effect will prevail over the intra-temporal if $\sigma > \theta$. Equation (8) depicts the relationship between consumption of tradables and non-tradables, with θ representing the elasticity of substitution among these goods. Finally, the equilibrium supply of non-tradables is presented in equation (9). Here, the higher is aggregate consumption, C, the lower is the level of production, as agents increase leisure along with the 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]

$$Y_N = C_N = \left(\frac{\theta-1}{\theta\kappa}\right)^{\frac{\sigma}{1+\sigma}}(1-\gamma)^{\frac{1}{1+\sigma}} \tag{10}$$

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

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

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

$$\tilde{Y}_N = \tilde{C}_N = \left(\frac{\sigma-\theta}{\sigma-1}\right)\gamma(\tilde{P}_N - \tilde{P}_T) + \left(\frac{\sigma}{\sigma+1}\right)\tilde{A}_N \tag{14}$$

Note that (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:

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

and its foreign counterpart is defined analogously,[7]

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

2.3 The real exchange rate equation

We define 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 (4) and (5) to define the log of the real exchange rate,

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

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

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

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. Deviations from the law of one price in tradables are large and persistent but stationary (Engel, 1993; Wei and Parsley, 1995), even in the presence of shipping costs (Obstfeld and Taylor, 1997). Hence, the unit root behavior in real exchange rates (q_t) might be induced by non-stationary behavior of y_t, which, in turn, 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:

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

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

III. Data and Empirical Implementation

3.1 The data

We pool a sample of 67 countries over the 1966-97 period to test the long-run equilibrium relationship between the real exchange rate and its determinants. Our dependent variable, the real exchange rate (q), is defined as 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:

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

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]

Data on net foreign assets (*NFA*) are drawn from Kraay, Loayza, Servén and Ventura (2000). Their 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:

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

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*. See Appendix I for the discussion of the methodological issues related to this database.

Domestic traded and non-traded productivity are proxied with the output per worker of the manufacturing and construction sector, respectively.[10] These data were drawn from the World Bank’s World Development Indicators and corrected for international comparison using the Summer-Heston’s PPP investment and output prices, respectively. Foreign productivity in both sectors is the trade-weighted average of productivity in the rest of the countries. 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 law of one price (LOP) follow a non-linear process in spatially separated market (Sercu and Uppal, 2000). Thus, x_{it} is mean reverting with the speed of adjustment varying directly with the extent of the deviation from LOP.

In practice, we assume that deviations from LOP in traded goods depend on transaction costs. Given the lack of cross-country data on transaction costs and relying on the empirically successful gravity equation model (Leamer and Levinsohn, 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 LOP in tradables might be approximated by country and time effects, that is:

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

or,

$$x_{it} = \eta_i + \mu_t \tag{21}$$

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

$$\tilde{y}_{it} = \left(\frac{\Psi}{1-n}\right) r\tilde{F}_{it} + \Psi(\tilde{Y}_T - \tilde{Y}_T^*)_{it} + \Psi(\tilde{P}_T^X - \tilde{P}_T^M)_{it} - \left(\frac{\sigma\Psi}{1+\sigma}\right)(\tilde{A}_N - \tilde{A}_N^*)_{it} + \xi_{it} \qquad (22)$$

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 equation (22) can also be expressed in log levels:

$$y_{it} = \frac{\Psi r}{(1-n)\gamma}\left(\frac{F}{Y}\right)_{it} + \Psi \ln\left(\frac{Y_T}{Y_T^*}\right)_{it} + \Psi \ln\left(\frac{P_T^X}{P_T^M}\right)_{it} - \left(\frac{\sigma\Psi}{1+\sigma}\right)\ln\left(\frac{A_N}{A_N^*}\right)_{it} + \xi_{it} \qquad (23)$$

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):

$$y_{it} = x_{it} + y_{it} = \eta_i + \mu_t + \beta_1\left(\frac{F}{Y}\right)_{it} + \beta_2 \ln\left(\frac{Y_T}{Y_T^*}\right)_{it} + \beta_3 \ln\left(\frac{P_T^X}{P_T^M}\right)_{it} + \beta_4 \ln\left(\frac{A_N}{A_N^*}\right)_{it} + \xi_{it} \quad (24)$$

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 thus they require a real exchange rate depreciation (*"transfer effect"*). Also, 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 $\beta_3>0$).

IV. Results

In order to consistently estimate the real exchange rate equation specified in (24), we use the recently developed panel cointegration techniques. Our empirical strategy consists of the following steps: (1) Testing for the presence of unit roots in the series involved in our analysis. (2) Testing for panel cointegration in the long-run real exchange rate equation. (3) Estimate the cointegrating relationship. In addition, we estimate our long-run real exchange rate equations for sub-samples of countries classified by income per capita and capital controls because of the heterogeneity of our sample.

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 apply the panel unit root tests implemented by Im, Pesaran, and Smith (1995). They 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 that averages all individual unit root test statistics.[11] 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.[12] IPS compute a $\bar{t}$-statistic, defined as the average of the individual *ADF-t* 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) and the critical values are tabulated by *Im*, Pesaran and Smith (1995).[13] Recently, Taylor and Sarno (1998) have proposed a multivariate test where the null hypothesis is that at least one of the series in the panel is a realization of a unit root process, whereas the alternative is that all of the series are realizations of stationary processes (i.e. H_1: $\rho_i < 1$, for all i). Their method applies Johansen's (1998) ML procedure of testing for the number of cointegrated vectors in a system. Although their alternative hypothesis is more rigorous towards stationarity than IPS's hypothesis, we cannot perform this test due to lack of sufficient time dimension T, given the number N of countries involved.

Before carrying out the ADF regressions for our panel unit root tests, we remove any common time effects (i.e. we regress our variables on a set of time dummies and save 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 difference-stationarity in our series, with the alternative implying 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. That is, all our series are integrated of order one, i.e. I(1).

4.2 Testing for cointegration

Our estimation results would be spurious if we do not test for panel cointegration in our model. We use the cointegration tests proposed by Kao (1999) and Pedroni (1995) to test whether equation (24) is valid in the long run. Kao (1999) has computed two type of panel cointegration tests: (i) DF-type tests (DF_ρ, DF_t, $DF_\rho{}^*$, $DF_t{}^*$) and, (ii) an ADF-type test. Regarding the DF-type tests, DF_ρ and DF_t are based on the strong exogeneity of regressors and errors, whereas $DF_\rho{}^*$ and $DF_t{}^*$ test for cointegration in the presence of endogenous regressors. The distribution of these statistics converges to a standard normal distribution.[14] On the other hand, Pedroni (1995) developed two sets of panel cointegration tests allowing for considerable heterogeneity. The first set involves averaging test statistics such as Phillips-Ouliaris tests for cointegration in time series across individuals. The second set performs averaging by pieces, so that the limiting distributions are based on limits of piece-wise numerator and denominator terms. The rejection of the null hypothesis in Pedroni's tests implies that enough of the individual cross-sections have statistics "far away'' from the theoretically predicted means if they were generated under the null.[15] Our panel cointegration results are presented in Table 2. There we find that all test statistics are significant so that the null of no cointegration is strongly rejected.

4.3 Correlation analysis

A first approximation to the comovement between real exchange rates and their fundamentals is undertaken with our correlation analysis. This analysis is conducted for a cross of 67 countries over 1966-97, for a panel data of nonoverlapping 8-year observations over the same period, and for panel data of annual information for the full sample and for different sub-samples. Given that our variables are non-stationary, we compute the correlation between the series in log differences. From the results reported in Appendix II, we find the following:

First, the real exchange rate, $\Delta ln\ q$, and net foreign assets, $\Delta(F/Y)$, 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, productivity in tradables, $\Delta ln(y_T/y_T^*)$, is positively and significantly correlated with $\Delta ln\ q$. This result is robust across the different groups of countries for both the cross-section and panel data sets. The cross-section correlation is stronger in low-income countries than in high-income countries (0.33 vs. 0.28), whereas the opposite result is found in panel data (0.18 vs. 0.42). In addition, the cross-section correlation between $\Delta ln(y_T/y_T^*)$, and $\Delta ln\, q$ 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, $\Delta ln\,(P_T^X / P_T^M)$, and $\Delta ln\, q$ 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 non-traded productivity, $\Delta ln(A_N/A_N^*)$, and $\Delta ln\, q$, that holds across sub-groups of countries for both cross-section and panel data sets. This positive unconditional correlation 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 $\Delta ln(A_N/A_N^*)$ and $\Delta ln\, q$ is negative for both the cross-section and panel correlation across countries, though statistically significant only in the latter.

4.4 Panel cointegration: the evidence

In this section we aim 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.

4.4.1 *On the estimation of panel cointegrated models: a very brief review*

For the estimation of the panel cointegrated systems, the literature uses the ordinary least squares (OLS), fully-modified least squares (FM), and dynamic least squares (DOLS) techniques. Studies have found that the limiting distributions of the FM and DOLS are asymptotically normal (Phillips and Moon, 1999; Kao and Chiang, 2000). 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.[16]

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 ($ln\ q_{it}$), 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 fundamentals. We assume that $\{x_{it}\}$~I(1), for all i,

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

In equations (26) and (27), we have Phillips' (1991) "*triangular representation*". Given the superiority of the DOLS estimator among the alternative techniques (Kao and Chiang, 1999), we decide to apply this technique to our real exchange rate equation.[17] The DOLS estimator, $\hat{\beta}_{\mathrm{DOLS}}$, can be obtained by running the following regression:

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

Details on the properties and limiting distribution of OLS and DOLS estimators are provided by Kao and Chiang (1999) and Phillips and Moon (1999). A summary of these statistical properties is presented by Calderón (2002).

4.4.2 *Estimation results*

In Tables 4 through 7 in Appendix II we present our estimation results for full sample of countries as well as for the different sub-samples of countries classified by income per capita and capital controls. Our preferred estimation technique is the dynamic least squares (DOLS) with 2 lags and 1 lead, that is, DOLS(2,1).

A. *Full sample of countries*

Our estimates for the full sample of countries are reported in Table 4 in Appendix II 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).

Although the OLS estimates are significant and have the expected signs, they are generally biased due to the presence of endogenous regressors. Hence, we present other estimates based upon OLS with bias correction, FM and DOLS.[18] Note that DOLS estimates are quite different from the FM, even though both estimators have the same limiting distribution (Kao and Chiang, 1999).

We first find that our coefficient estimates have the expected sign and are statistically significant for the 1966-97 period, regardless of the estimation technique. Second, we find that our OLS estimates are downward biased. When comparing our DOLS and OLS estimates, we 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.

It seems very restrictive to assume *"country homogeneity"* in our estimates for the long-run real exchange rate equation. The effect of the different fundamentals on the real exchange rates could differ among countries due to different levels of income or the imposition of capital controls. For this reason, we proceed to estimate equation (24) for different sub-samples of countries. Table 5 in Appendix II 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 in Appendix II we present estimates for subsample of countries according to the presence of capital controls and the degree of black market premium in the foreign exchange market, respectively.[19] 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*

We first test the robustness of our long-run estimates by adjusting our real exchange rate equation to different sub-samples of countries classified by income levels and capital controls. In the former case, we use World Bank's classification of countries by income per capita,[20] 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 use two different criteria: (a) we define the presence of capital controls by aggregating a set 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 multiple exchange rate practices, current account restrictions, capital account restrictions, and 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. (b) 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 net foreign assets for the full sample of countries as well as for the samples of high- and low-income countries. The effect of F/Y is larger for low and lower-middle income countries being larger (0.36 vs. 0.18), thus denoting a more powerful *transfer effect* for this group of 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).[21]

Second, surges in relative productivity in tradables will generate a real exchange rate appreciation. This result is robust whether we use different indicators of relative productivity (relative output per worker or relative total factor productivity) or different time periods. It holds not only for the full sample but also for all the sub-samples presented from Tables 5 through 7 in Appendix II. Note that the coefficient of $ln\,(Y_T/Y_T^*)$ is higher for high-income countries than for low-income countries. 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, a decline in the terms of trade may be associated with real exchange rate depreciations. This positive relationship holds for the full sample and for sub-samples of both high and low-income countries, although it is not significant for latter sub-sample. On the other hand, 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 0.7 percent depreciation in the real exchange rate.

Fourth, positive shocks in non-traded productivity may depreciate the real exchange rate. We find that $ln\,(A_N/A_N^*)$ has a negative and significant coefficient for the full sample of countries and for the sample high-income countries. In addition, this finding holds for the samples of countries with low capital controls and low black market premium. 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 0.2 percent depreciation in the real exchange rate.

V. 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 in Appendix II we present formal tests of homogeneity between and within groups of countries. First, we test the equality of the coefficient estimates for 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 (individual and joint) Hausman-type tests in the spirit of Pesaran, Shin and Smith (1999). Here we test whether our panel DOLS estimates are statistically equal to an average of DOLS estimates performed on a country-by-country basis (see Tables 8-10 for more details).[22]

A. *Testing the equality of coefficients across sub-groups of countries*

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

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 of net foreign assets, F/Y, are statistically different across groups. See bottom panel of Tables 6 and 7 in Appendix II.

B. *Testing the hypothesis of country homogeneity*

Here we test for country homogeneity behind the long-run estimates presented in Tables 4-7 in Appendix II, that is, we test whether the *'pooling assumption'* is valid for our long-run estimates in the sample of all countries (Table 8 in Appendix II) and in all sub-samples (Tables 9 and 10 in Appendix II).

First, we find that our pooled and average DOLS estimates of *F/Y* and $ln\,(A_N/A_N^*)$ are statistically equal for the full sample of countries over 1966-97, whereas the coefficients of $ln\,(P_X/P_M)$ and $ln\,(y_T/y_T^*)$ are not statistically equal across estimators (see Table 8). The joint Hausman test rejects the null of homogeneity in the limit (p – value = 0.059). In addition, 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 for country heterogeneity in the equations for high- and low-income countries (see Table 9 in Appendix II). We find that the null 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 *F/Y*) favored the hypothesis of heterogeneity.

Third, we test the pooling assumption for the sub-samples of countries with low and high capital controls (see Table 10 in the Appendix). Here we find that both individual and joint Hausman tests fail to reject the null of homogeneity for countries with low capital controls, whereas the opposite result is found for countries with high capital controls (i.e. we reject the null of homogeneity). 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

After testing parameter constancy across countries, we test the structural stability of our cointegrating relationship over time. There is a wide array of papers on structural change testing for non-stationary time series. However, recent research has extended this analysis to non-stationary panel data.

We use the methodology developed by Kao and Chiang to test the stability of our cointegrating vector over time.[23] They propose Wald-type test statistics for changes in the cointegrating vector at unknown break points. Also, they find that the limiting distribution of this Wald test is free of nuisance parameters (i.e. the square of a Bessel process), and similar but not identical to that of the test developed by Andrews (1993). Finally, their test has non-trivial power against a wide array of alternatives, that is, 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).[24] 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 in Appendix II). 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

To assess the plausibility of our long-run coefficient estimates, we compare them with the calibrated parameters obtained if we use the technology and preferences parameter values in the business cycle literature for industrial countries (Stockman and Tesar, 1995) and for developing countries (Mendoza, 1991). 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. For simplicity, we assume that developing countries would only differ in the elasticity of intertemporal substitution, which takes the value of 0.999. Finally, the country size of the representative high-income 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 equation.

The coefficient estimates for the full sample (see regressions 1 and 3 in Table 4 in Appendix II) are broadly in line with the calibrated parameters. In regression 1 (1966-97), 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 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 in Table 4 in Appendix II). If we reduce the elasticity of intertemporal substitution (σ) such that the coefficient estimate for the relative traded-sector productivity, $ln\ (Y_T/Y_T^*)$, and the calibrated parameter are equal (setting $\sigma = 0.294$, and making $\Psi = 1.3024$), we find that the estimate of *F/Y* is closer to the calibrated parameter (0.2127 vs. 0.1792, respectively). On the other hand, the coefficient of $ln\ (P_X/P_M)$ is not equal to the coefficient of $ln\ (Y_T/Y_T^*)$, being smaller than the calibrated parameter (0.7427 vs. 1.3037). Finally, the coefficient of $ln\ (A_N/A_N^*)$ is also smaller (in absolute value) than the calibrated value (–0.1856 vs. –0.2961).

On the other hand, discrepancies between estimates and calibrated parameters for the sample of high-income countries are slightly larger than the ones obtained for the full sample of countries. Using the results from regression 1 in Table 5, we find that the estimated coefficient of *F/Y* doubles the value of the calibrated parameter (0.1804 vs. 0.0830). The coefficient of $ln\ (P_X/P_M)$ is slightly larger than the calibrated parameter (1.3649 vs. 1.0870), whereas the coefficient of $ln\ (A_N/A_N^*)$ is four times larger than the calibrated parameter (–1.4078 vs. –0.3623). Then, we modify σ such that the estimated and calibrated coefficient for $ln\ (P_X/P_M)$ are equal (i.e. $\sigma = 0.2715$ and $\Psi = 1.3649$). We find that the estimated coefficients for *F/Y* and the $ln\ (A_N/A_N^*)$ are closer to the calibrated parameters (0.1021 and –0.2857, respectively), however, we still have different coefficient estimates for $ln\ (P_X/P_M)$ and $ln\ (Y_T/Y_T^*)$.

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

After assessing the long-run behavior of real exchange rates, we proceed to analyze its short-run dynamics. Although we constrained the long-run coefficients to be equal across countries,[25] 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 different monetary regimes.

In practice, we take the cointegrating vector estimated for the full sample of countries and we estimate an 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 an implied half-life of these deviations equal to $ln(0.5)/ln(1+\phi_i)$. We denote ECM_1 as the short-run model specified in (29). On the other hand, we 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.[26] Hence, the alternative short-run model is denoted by ECM_2,

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

Given that $\Delta(m-m^*)$ helps 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 average and 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. From our country estimates of $\hat{\phi}_i$, 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 monetary 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 Figures 1 and 2).

FIGURE 1
AVERAGE HALF-LIFE OF REAL EXCHANGE RATE DEVIATIONS FROM THE EQUILIBRIUM, 1973-97

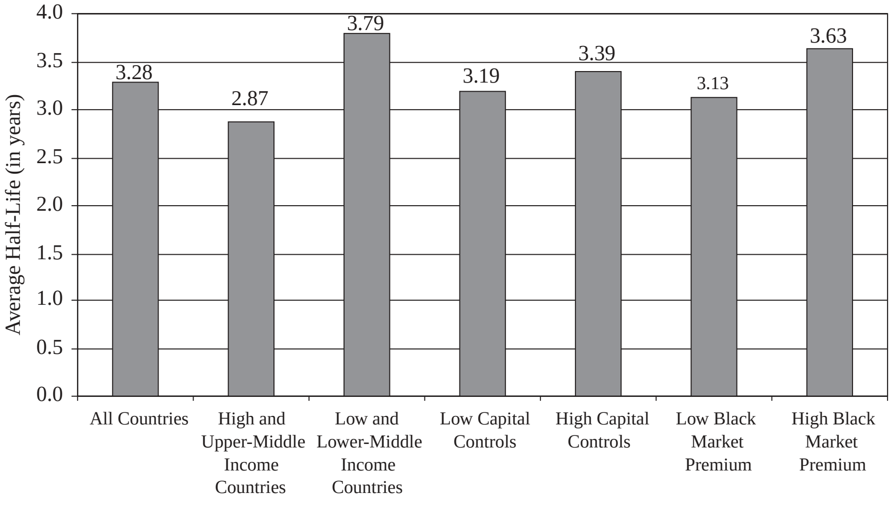
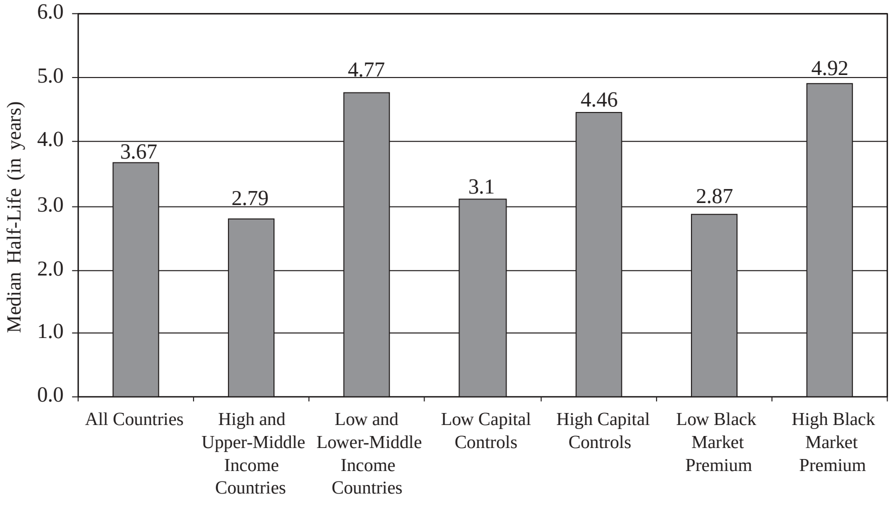


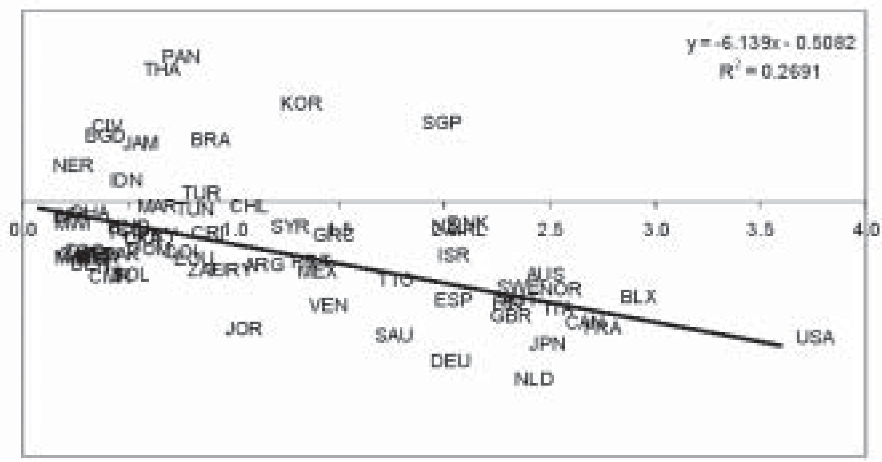
FIGURE 2
MEDIAN HALF-LIFE OF REAL EXCHANGE RATE DEVIATIONS FROM THE EQUILIBRIUM, 1973-97



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.

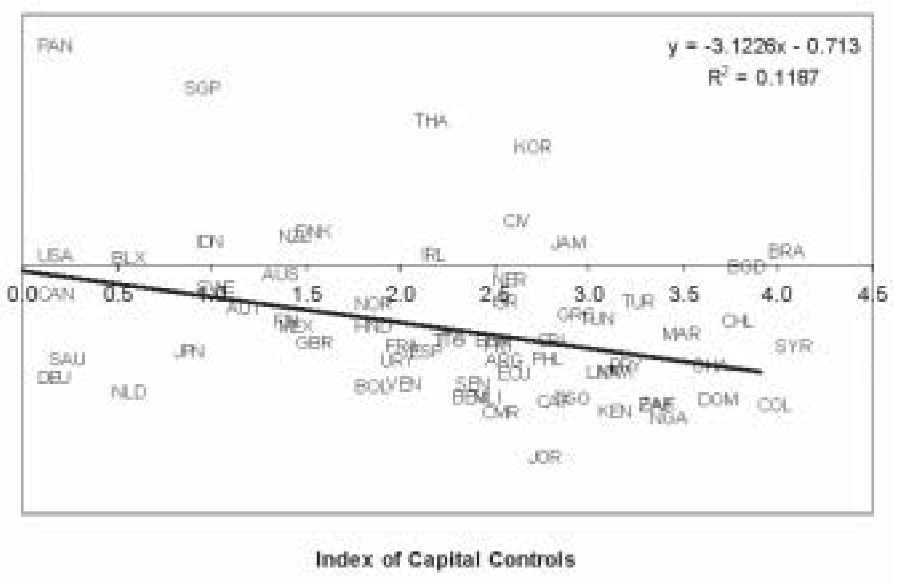
Finally, we roughly characterize the half-life of real exchange rate deviations from the equilibrium by evaluating their link with output per worker and capital controls. We find that the half-life of deviations from the equilibrium exchange rate is negatively associated with income per worker (–0.13), 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).

FIGURE 3
CONDITIONAL CORRELATION BETWEEN THE RATIO OF INCOME PER WORKER AND THE HALF-LIFE OF REAL EXCHANGE RATE DEVIATIONS



Ratio of Income per Worker (Home vs. Rest of the World)

FIGURE 4
CONDITIONAL CORRELATION BETWEEN THE INDEX OF CAPITAL CONTROLS AND THE HALF-LIFE OF REAL EXCHANGE RATE DEVIATIONS



VII. 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 (i.e. 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 (i.e. productivity shocks, demand shocks) could be achieved with the implementation of structural vector autoregression models.

Notes

1 Lane and Milesi-Ferreti (2000) have followed 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.

2 Disutility in producing more output is captured by the term $-(\kappa/2)\,y^2_{N,s}$. Assuming that disutility from effort ℓ_N is given by $-\psi\ \ell_N$ and that $y_N = A\ell_N^{\alpha}$ $(\alpha < 1)$, then $\kappa = 2\,\psi\,/\,A^{1/\alpha}$. The output term in equation (1) is obtained when $\alpha = 0.5$. A rise in productivity A is here captured by a fall in κ (Obstfeld and Rogoff, 1996).

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.

4 In addition to the first-order conditions (7)-(9), and the period budget constraint, equation (3), we require the transversality condition: $\lim_{T\to\infty} \prod_{v=s}^{t+T}\left(\frac{1}{1+r_v}\right)F_{t+T+1}$ to characterize our equilibrium.

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

6 In steady state we find that: (a) if work effort is less taxing (smaller κ), production of non-tradables will be larger (equation 10); (b) if consumption of tradables has more weight in the utility function (larger γ), the ratio of traded to non-traded output is higher.

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

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 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.

10 We also run our regressions with non-traded sector productivity as labor productivity in the services sector and the results were qualitatively invariant. However, given the increasing trade in services we prefer to use output and labor in construction as proxies for non-traded labor productivity.

11 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.

[12] 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.

[13] It 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 (Choi, 1999). However, linear time trends are included in the model, its power decreases considerably (Breitung, 1999; Choi, 2001).

[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.

[16] 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.

[17] This estimator is a panel version of the dynamic least squares estimator proposed by Phillips and Loretan (1991), Saikkonen (1991), and Stock and Watson (1993).

[18] The FM estimation corrects the dependent variable using the long-run covariance matrices to remove nuisance parameters and applies OLS to the corrected variables.

[19] For a complete list of countries and its classification according to income level or capital controls, see Table 13 in Appendix II of results.

[20] 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.

[21] 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 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.

[22] 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.

[23] For further details on this type of testing, see Kao and Chiang (2000). Also, Calderón (2002) presents a summary of this test in the Appendix C.

[24] Although not reported in Table 11, 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.

[25] 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).

[26] We presume that by including the monetary variable, the speed of adjustment to the long run would be faster.

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APPENDIX I

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).*

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

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 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.** From these initial values, stock series are obtained using the recursive formula:

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

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

** 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. This is not too serious a problem given that portfolio flows are a relatively recent phenomenon. Hence, in the absence of 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.

weights given by the structure of intra-OECD flows; we set δ at 4 percent. For portfolio equity liabilities, we set $\delta = 0$ 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.

APPENDIX II

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	1966-97	1973-97	1973-97
	Using y_T/y_T^*	Using A_T/A_T^*	Using y_T/y_T^*	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					
$\Delta(F/Y)$	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)					
$Dln(A_N/A_N^*)$	–0.0288	–0.0346	0.0859	0.1705	–0.1636
II. Panel Data Correlations					
$\Delta(F/Y)$	–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
$\Delta(F/Y)$	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 **

Notes: 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	1973-97	1966-97	1973-97
	Using y_T/y^*_T	Using y_T/y^*_T	Using A_T/A^*_T	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

Table 4 (Cont.)

Variables	1966-97	1973-97	1966-97	1973-97
	Using y_T/y_T^*	Using y_T/y_T^*	Using A_T/A_T^*	Using A_T/A_T^*
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

Variables	1966-97		1973-97	
	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	1966-97	1973-97	1973-97
	FM-OLS	DOLS	FM-OLS	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

Variables	1966-97		1973-97	
	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 Capital 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) **	4.267 (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]
Variables	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 (ϕ_i) AND HALF-LIFE OF EQUILIBRIUM DEVIATIONS (h_i)
Sample Period: 1973-97

Region	Statistic	Model ECM$_1$		Model ECM$_2$	
		ϕ_i	h_i	ϕ_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 ϕ, and the half life of the deviations from the equilibrium, $h = \ln(0.5)/\ln(1+\phi)$.

TABLE 13
CLASSIFICATION OF COUNTRIES

Country	Classification by Income Levels		Classification by Capital Controls			
	High and Upper-Middle	Low and Lower-Middle	Low Capital Restrictions	High Capital Restrictions	Low Black Market Premium	High Black Market Premium
Argentina	X		X		X	
Australia	X		X			X
Austria	X		X			X
Benin		X	X			X
Burkina Faso		X	X			X
Bangladesh		X		X	X	
Belgium-Luxemburg	X		X			X
Bolivia		X	X		X	
Brazil	X			X	X	
Central African Republic		X		X		X
Canada	X		X			X
Chile	X			X	X	
Côte d'Ivoire		X	X			X
Cameroon		X	X			X
Colombia		X		X		X
Costa Rica		X		X	X	
Germany	X		X			X
Denmark	X		X			X
Dominican Republic		X		X	X	
Ecuador		X		X	X	
Spain	X		X			X
Finland	X		X			X
France	X		X			X
United Kingdom	X		X			X
Ghana		X		X	X	
Greece	X			X		X
Honduras		X	X			X
Indonesia		X	X		X	
India		X		X	X	
Ireland	X		X			X
Israel	X			X		X
Italy	X		X			X
Jamaica		X		X		X
Jordan		X		X		X
Japan	X		X			X
Kenya		X		X		X
Korea	X			X		X
Sri Lanka		X		X	X	

Table 13 (Cont.)

Country	Classification by Income Levels		Classification by Capital Controls			
	High and Upper-Middle	Low and Lower-Middle	Low Capital Restrictions	High Capital Restrictions	Low Black Market Premium	High Black Market Premium
Morocco		X		X		X
Mexico	X		X			X
Mali		X		X		X
Malawi		X		X	X	
Niger		X	X			X
Nigeria		X		X	X	
Netherlands	X		X			X
Norway	X		X			X
New Zealand	X		X			X
Pakistan		X		X	X	
Panama	X		X			X
Peru		X	X		X	
Philippines		X		X		X
Portugal	X		X			X
Paraguay		X		X	X	
Saudi Arabia	X		X			X
Senegal		X	X			X
Singapore	X		X			X
Sweden	X		X			X
Syrian, Arab Rep.		X		X	X	
Togo		X		X		X
Thailand		X	X			X
Trinidad and Tobago	X		X		X	
Tunisia		X		X	X	
Turkey	X			X		X
Uruguay	X		X			X
United States	X		X			X
Venezuela	X			X	X	
South Africa		X		X		X