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Prakesh Apte, Piet Sercu and Raman Uppal

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Prakesh Apte, IIM Bangalore
Piet Sercu, Katholieke Universiteit Leuven
Raman Uppal, London Business School (LBS) and CEPR

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Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR, UK
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: www.cepr.org

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ABSTRACT

The Exchange Rate and Purchasing Power Parity: Extending the Theory and Tests*

This Paper analyses the exchange rate in a 'no-arbitrage' or 'real business cycle' equilibrium model and provides empirical evidence for this model *vis-a-vis* PPP. Our contribution is to show, based on a generalization of the equilibrium model of exchange rates, that (i) the test equation linking the exchange rate to fundamentals should allow for international heterogeneity in time preferences or risk attitudes, as well as noise – that is, the model should not be tested as an exact relation; (ii) empirical work should use levels of variables rather than first differences; (iii) tests on the existence of long-run relations should be complemented by tests on the signs of the coefficients; (iv) the specification of the regression should offer demonstrated advantages over alternatives, and the significance tests should not rely on asymptotic distributions; and (v) the tests should steer clear of countries that have imposed, for most of the period, capital restrictions or exchange controls, thus violating the integrated-markets assumption of the model. Our empirical work shows that, as a long-run relation, the generalized model outperforms PPP.

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Prakesh Apte
IIM Bangalore
Bannerghatta Road
Bangalore
560 076 INDIA
Email: apte@iimb.ernet.in

Piet Sercu
Centrum voor Economische Studien
Katholieke Universiteit Leuven
Dekestraat 2
Naamsestraat 69
B-3000 Leuven
BELGIUM
Tel: (32 16) 28 57 72
Fax: (32 16) 28 57 99
Email: piet.sercu@econ.kuleuven.ac.be

Raman Uppal
Institute of Finance and Accounting
London Business School
Regents Park
London
NW1 4SA
Tel: (44 20) 7706 6883
Fax: (44 20) 7724 3317
Email: ruppal@london.edu

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1 Introduction

A recent class of exchange rate models (called by a variety of names such as "equilibrium", "no-arbitrage", and "real business cycle" models) relates the real exchange rate to real consumption, time preference, and relative risk aversion (or intertemporal substitution). Our main objective, in this paper is to extend the existing models, and add fresh empirical evidence to extant tests of this model. This introductory section provides a brief review of the theoretical foundations and available empirical results.

The cornerstone of the equilibrium approach to exchange rates is the result that, under suitable assumptions, the real exchange rate is the ratio of the marginal utilities of real spending.[1] For instance, Stockman (1980) and Lucas (1982) study the exchange rate in a model of an endowment economy with frictionless markets, where the solution is of an "equilibrium" nature in the sense that it maximizes international aggregate utility. Dumas (1992) applies this equilibrium property of the exchange rate in his analysis of a one-good production economy where international trade in goods is costly[2] and where agents have constant and identical risk aversions and time preferences. Sercu, Uppal, and Van Hulle (1995) study an endowment-economy with shipping costs and obtain closed-form solutions rather than the numerical results of Dumas. Instead of using shipping costs to model countries as distinct economic entities, Backus and Smith (1993) derive the exchange-rate in a model with one perfectly tradable good and one non-tradable good and CES consumption preferences defined over these two goods.

Other utility-maximizing models often do not refer to this literature, but nevertheless fit into the "equilibrium" category—Stulz (1987), for example, considers a two-country production economy with log investors who have a identical Cobb-Douglas preferences defined over a perfectly tradable good and a non-traded good. The equilibrium model also nests the monetary economies considered in Bakshi and Chen (1997) and Basak and Gallmeyer (1999), where the focus is on determining the prices of financial securities rather than expressing the exchange rate in terms of observable variables and relating it to PPP. Bakshi

[1] For a recent review of macroeconomic models of the exchange rate see Devereux (1997), Lane (1999), and Obstfeld and Rogoff (2000).

[2] Empirical support for the effects of shipment costs has been documented in Engel (1993), Rogers and Jenkins (1995), and Wei and Parsley (1995), who find that a significant proportion of the total variation in the real exchange rate arises from deviations from the Law of One Price (LOP). Also, Engel and Rogers (1995) show that within-country deviations from LOP are much smaller than cross-country deviations.

and Chen assume log utilities, lognormal distributions, and perfect commodity markets where money is neutral, while Basak and Gallmeyer characterize the exchange rate in an exchange economy with money in the utility function. In all these models, market failures are absent, so that the solution corresponds to (and is often explicitly based on) the central-planners solution.

Backus, Foresi and Telmer (1996) generalize the Stockman (1980) and Lucas (1982) models rather than specializing it. First, they link the *nominal* exchange rate to the marginal utility of *nominal* spending, an approach that neatly avoids the assumption, implicit in any real-exchange-rate model with more than one good, that utility is homothetic. Moreover, the Backus *et al.* result is established starting from the notion of no arbitrage in financial markets, where the key assumption is that asset markets are frictionless and complete; thus, the "equilibrium" model is robust to market failures in the real economy provided they do not invalidate the role of marginal utility as the pricing kernel.[3] Sercu and Uppal (2000) provide a general first-order decomposition of the change of the nominal exchange rate for a wide range of (state-independent) utility functions and economic settings; they also discuss the links of the equilibrium model with the PPP hypothesis. Our contribution on the theoretical front is to extend the Sercu and Uppal (2000) model to allow for state dependent utility functions, and more importantly, different rates of time-preference and risk aversion across countries and across time.

Turning now to empirical work, we first review test of the PPP hypothesis. Abuaf and Jorion (1990) use Dickey-Fuller tests to establish the presence of mean-reversion in real exchange rates, a phenomenon not evident from autocorrelations tests on first-differenced data; and Johansen and Juselius (1992) refine the cointegration techniques in an attempt to find long-run relations between exchange rates and relative price levels that may be hard to detect in first-differenced data. Edison, Gagnon, and Melick (1997) show how the power of these tests can be improved using the Horvath and Watson (1995) procedure. Expanded data sets have been considered by, for instance, Frankel and Rose (1996), Froot, Kim and Rogoff (1995), Lothian (1997), Lothian and Taylor (1995), Wei and Parsley (1995), and Taylor (1995). Related work includes Engel, Hendrickson and Rogers (1996) and O'Connell (1998). A review of the empirical literature on PPP is provided by Froot and Rogoff (1995),

[3]See Duffie (1992) for details on this approach to asset pricing. The pricing-kernel or martingale-pricing approach has also been used to study the relation between exchange rates and international interest rates by Nielsen and Saa-Requejo (1993), Backus, Foresi and Telmer (1996), and Hollifield and Uppal (1997).

Nessén (1994) and Rose (1996/7).

We now review the empirical tests of the equilibrium model of exchange rates. Backus and Smith (1993) test (and reject) the hypothesis that the real exchange rate is an exact loglinear function of the real-consumption ratio, by comparing the means, standard deviations, and autocorrelations of changes in the two time series. Most tests, however, rely on regression analysis, either of differenced data or of levels. An example of the first type of regression is found in the first-pass tests by Koedijk Nissen, Schotman and Wolff (1996), where they cross-sectionally relate changes in exchange rates to consumption growth rates and find no significant link. When subsequently studying panel data at the integrated level, Koedijk *et al.* find evidence broadly consistent with PPP but, again, no traces of any consumption effect. Other tests that analyze levels of the data are cointegration tests as provided by, for instance, Nessén (1994) and Sercu and Uppal (2000). These tests typically find evidence in support of the existence of a long-term relation between nominal exchange rates and CPIs, but with a cointegration vector that is incompatible with PPP. In addition, in the analysis of quarterly data on five countries by Sercu and Uppal (2000) there appears to be a long-run link between the real exchange rate and real consumptions, which seems to favor the "equilibrium" model over PPP. However, cointegration tests by Kollmann (1995) on seven countries do not find evidence of such relations.[4]

In Section 2, we start from these empirical studies to come up with a list of desiderata for our own work: (i) at the very least, the test equation should allow for international heterogeneity in time preferences or risk attitudes, as well as noise—that is, the model should not be tested as an exact relation; (ii) empirical work should use levels of variables rather than first differences; (iii) tests on the existence of long-run relations should be complemented by tests on the signs of the coefficients; (iv) the specification of the regression should offer demonstrated advantages over alternatives, and the significance tests should not rely on asymptotic distributions; and (v) the tests should steer clear of countries that have imposed, for most of the period, capital restrictions or exchange controls, thus violating the integrated-markets assumption of the model. Our empirical work shows that, as a long-run relation, the generalized model does outperform PPP and that, with one exception, the coefficients have the correct sign. Section 3 describes our data and test procedure. Section 4 discusses the empirical results. We conclude in Section 5.

[4]Apart from the base country, the US, the only country that shows up in both papers is Japan. We return to the sample selection issue in Section 2.

2 The model and test design

This section presents the considerations motivating the design of our test of the RBC model of exchange rates. Let $S(t)$ denote the nominal exchange rate (units of currency 1 per unit of currency 2). Backus *et al.* (1996) establish that, in an economy with representative consumers and perfect, arbitrage-free capital markets, the nominal exchange rate equals the ratio of the marginal utilities of nominal spending of country 2 and country 1. That is , if country k's marginal utility is denoted by $m_k(t)$, $k = \{1, 2\}$, then

$$S(t) = \frac{m_2(t)}{m_1(t)}. \tag{1}$$

Let $C_k(t)$ denote nominal consumption in country k at time t, η the measure of relative risk aversion of the representative investor, and $\Pi_k(t)$ the consumption price level in country k at time t. If utility is iso-elastic with exponent $(1 - \eta)$, then the marginal utility of *real* consumption is $(C(t)/\Pi(t))^{-\eta}$, and the utility of nominal spending is obtained by dividing by $\Pi(t)$. Rearranging, we obtain the model for the real rate that is tested by Backus and Smith (1992) and Kollmann (1995):

$$S(t)\frac{\Pi_2(t)}{\Pi_1(t)} = \left(\frac{C_1(t)/\Pi_1(t)}{C_2(t)/\Pi_2(t)}\right)^{\eta}. \tag{2}$$

Given that the above model links the real rate to the consumption ratio, it is sometimes called the real business cycle (RBC) model. Being more precise than the labels "equilibrium" and "no arbitrage," we adopt the RBC name in the remainder of the paper. Note that the prediction is that higher real consumption abroad *lowers* the real value of the foreign currency. For testing purposes, the standard RBC model is often written in terms of percentage changes or changes of logs, see for instance Backus and Smith (1992) or Kollmann (1995):

$$\frac{dS}{S} = \eta\left[\left(\frac{dC_1}{C_1} - \frac{d\Pi_1}{\Pi_1}\right) - \left(\frac{dC_2}{C_2} - \frac{d\Pi_2}{\Pi_2}\right)\right] + \frac{d\Pi_1}{\Pi_1} - \frac{d\Pi_2}{\Pi_2}. \tag{3}$$

To see how much generality has been lost in the standard RBC equations (2)–(3), we now consider a setting that is less restrictive than that considered in the existing literature. In general, the marginal utility of nominal spending in country k at date t is an indirect

one, derived from the (static) problem of a consumer who faces a vector of prices for the N goods, $p_k(t)$, and who wishes to allocate a budget of $C_k(t)$ over the consumption of these goods, $c_k(t) \equiv \{c_{k1}(t), c_{k2}(t), \ldots, c_{kN}(t)\}$, in order to maximize utility, $U_k(c_k(t), X_k(t), t)$, where $X_k(t)$ is a vector of (possibly country-specific) state variables that affect utility.[5] This problem can be written as:

$$V(C_k(t), p_k(t), X_k(t), t) \equiv \max_{c_{kj}(t)} \left\{ (U_k(c_k(t), X_k(t), t) - \Lambda_k \left[\sum_{j=1}^{N} c_{kj}(t) p_{kj}(t) - C_k(t) \right] \right\},$$

where

$V(C_k(t), p_k(t), X(t), t)$ refers to the period-t indirect utility function of total spending, given prices,

$C_k(t)$ denotes the nominal consumption budget, expressed in terms of country k's currency,

$p_k(t)$ denotes the vector of local-currency price of good j in country k, $p_{kj}(t)$,

$X_k(t)$ denotes the vector of state variables that affect utility in country k,

$U_k(c_k(t), X_k(t), t)$ denotes the utility function of the representative investor in country k, and implicitly includes the discounting for time,

$c_k(t)$ denotes the vector of consumption quantities $c_{kj}(t)$ of good $j = \{1, \ldots, N\}$ consumed by the representative individual in country $k = \{1, \ldots, K\}$ at time t.

Thus, the marginal indirect utility of nominal spending in country k is the multiplier in the above optimization problem:

$$m_k(t) = \Lambda_k(C_k(t), p_k(t), X_k(t), t) = \frac{\partial V(C_k(t), p_k(t), X_k(t), t)}{\partial C_k(t)}. \tag{4}$$

The change in marginal utility can therefore be decomposed into a (possibly varying) time-preference component, growth in real consumption weighted by (possibly varying) relative risk aversion, an inflation rate computed from marginal spending weights, and fluctuations due to state variables. The details of this decomposition, a generalization of a result in

[5] The optimal level of $C_k(t)$, itself, would be obtained by solving the intertemporal problem of the consumer.

Sercu and Uppal (2000), is provided in the appendix. It then follows from equation (1) that the first-order components of the change in the exchange rate are given by

$$\begin{aligned}\frac{dS}{S} &= \underbrace{[\delta_2(t)-\delta_1(t)]\,dt}_{\text{differential time preference}} + \underbrace{\eta_1(t)\left(\frac{dC_1}{C_1}-\frac{d\Pi_1}{\Pi_1}\right)-\eta_2(t)\left(\frac{dC_2}{C_2}-\frac{d\Pi_2}{\Pi_2}\right)}_{\text{differential RRA-weighted real consumption growth}} \\ &+ \underbrace{\frac{d\pi_1}{\pi_1}-\frac{d\pi_2}{\pi_2}}_{\text{differential marginal inflation}} + \underbrace{\sum_{s=1}^{M}\zeta_{1s}\frac{dX_{1s}}{X_{1s}}-\sum_{s=1}^{M}\zeta_{2s}\frac{dX_{2s}}{X_{2s}}}_{\text{effects of state variables}} \qquad (5)\end{aligned}$$

where

$$\delta_k(t) \equiv -\frac{1}{\partial V_k/\partial C_k}\frac{\partial^2 V_k}{\partial C_k \partial t},$$ the semi-elasticity of marginal utility with respect to time, that is, the measure of instantaneous time preference,

$$\eta_k(t) \equiv -\frac{C_k}{\partial V_k/\partial C_k}\frac{\partial^2 V_k}{\partial C_k^2},$$ the degree of relative risk aversion,[6]

$$\frac{d\Pi_k}{\Pi_k} \equiv \sum_{j=1}^{N}\left(\frac{c_{kj}p_{kj}}{C_k}\right)\frac{dp_{kj}}{p_{kj}},$$ inflation weighted on the basis of total consumption,[7] and

$$\frac{d\pi_k}{\pi_k} \equiv \sum_{j=1}^{N}\left(\frac{\partial c_{kj}}{\partial C_k}p_{kj}\right)\frac{dp_{kj}}{p_{kj}},$$ inflation weighted on the basis of marginal consumption,[8]

$$\zeta_{ks}(t) \equiv \frac{X_{ks}}{\partial V_k/\partial C_k}\frac{\partial^2 V_k}{\partial C_k \partial X_{ks}},$$ the elasticity of marginal utility with respect to X_{ks}.

Thus, the standard RBC equations (2)–(3) assume that utility is homothetic as well as state-dependent (implying, respectively, $d\Pi_k(t)/\Pi_k(t) = d\pi_k(t)/\pi_k(t)$) and $\zeta_{ks}(t) = 0$). In addition, investors are assumed to have equal time preferences δ_k and constant and equal relative risk aversions η_k. A first obvious implication is that, if the primary issue of interest

[6] This definition of relative risk aversion, also adopted by Breeden (1978), is a "real" measure of relative risk aversion because, when taking partial derivatives with respect to C_k, we hold prices constant. In the one-good no-inflation case, this definition is identical to the standard definition, $-c_k[\partial^2 U_k/dc_k^2]/[\partial U_k/dc_k]$.

[7] When money is in the utility function, as in Stulz (1987), the interest rate will be part of the price index, Π_k.

[8] The marginal weights, $[\partial c_{kj}/\partial C_k]p_{kj}$, sum to unity by virtue of the budget constraint.

is the validity of the general approach rather than a highly specialized version of it, then time preferences and risk aversions ought not be restricted to be equal across time, as done in Backus and Smith (1992), Kollmann (1995) and Koedijk *et al.* (1996). That is, the RCB model should be extended to allow for a trend as well as for differential sensitivity to the two cycles.

Second, consider the implications of ignoring the difference between marginal- versus average-weighted inflation. One effect is that this introduces an error-in-variables bias, (reinforcing a similar problem arising from having to use noisy aggregate consumption data). But the deviation between the marginal and average inflation rates also implies that the relation between the exchange rate, consumption data, and CPIs is no longer exact even if risk aversions and time preferences were really constant and identical, and if the assumptions of complete markets, consensus probabilities and representative consumers were actually true. Other potential sources of inexactness that are assumed away in equation (2) are the variability in the cross-country time-preference differential and other state variables. At best, these omitted variables add noise to the equation, but if they are correlated with the regressors then they also induce an omitted-variables bias in the estimates of δ and η. While there are no obvious cures for the errors-in-variables and omitted-variables biases, the least one can do, in light of equation (5), is not just to extend the RBC model so as to allow for international heterogeneity in the δs and ηs, but also to allow for the presence of an error term. This error term in the constant-parameter regression will be autocorrelated if the risk- and time-attitude parameters change slowly over time.[9]

Koedijk *et al.* (1996) complement their first-pass regressions on differenced data by a thorough analysis of integrated data, which offers more power provided that there is a long-run relation between the levels of the variables. Such cointegration tests are provided by, amongst others, Kollmann (1995) and Sercu and Uppal (2000). While Kollmann finds no relation, Sercu and Uppal do. However, Sercu and Uppal do not establish whether the cointegration vector is consistent with PPP. In the same vein, a long-run relation between real exchange rate and real consumption data could still be incompatible with the RBC model if the real rate were negatively (positively) related to domestic (foreign) real consumption. That is, while, the generalized RBC model does not give specific numerical values for the η

[9]Note also that, in the empirical PPP literature, persistent deviations have never been viewed as sufficient grounds for rejection of PPP, even though in the PPP case there a no grounds to predict slowly moving parameters.

coefficients, we do know that relative risk aversion should be positive. Thus, we complement the available cointegration tests by estimating the magnitude and testing the sign of the coefficients (obtained from integrated data).

The new issue that arises immediately is about the regression specification and significance level to use. Koedijk *et al.* chose a specification that corresponds to the standard VAR/ECM equation for first differences of $S(t)$ except that also the contemporaneous inflation rates show up on the right hand side. They do not discuss any alternatives. A paper that does so, at great depth, is by Phillips and Loretan (1991), that sets up extensive Monte-Carlo experiments for various possible specifications of the equation. The one that performs the best contains the regressor as an integrated variable rather than in first-differenced form. Suppose, for simplicity of notation, that one wishes to study a relation $Y(t) = AX(t)$. The Phillips-Loretan equations would then be non-linear relations of the type

$$\begin{aligned} Y(t) &= AX(t) + a_{+1}[Y(t+1) - AX(t+1)] + a_{-1}[Y(t-1) - AX(t-1)] \\ &\quad + b_{-1}[X(t) - X(t-1)] + c_{-1}[Y(t) - Y(t-1)] + e(t) \end{aligned} \tag{6}$$

with additional leads and lags if required on empirical or *a priori* grounds. Thus, we adopt this form for our test equation. Phillips and Loretan also show that, especially in very large samples, with this regression specification the standard inferences from the usual t-tests are not flagrantly misleading. In our case, however, the sample (87 quarterly observations) is far from large. Thus, our inferences are based on extensive Monte-Carlo experiments set up under three alternative versions of the null of no relation between real rate and real spending. The details are provided in the next section.

Having chosen the empirical test specification, the next step is to select the data series. Both Backus and Smith (1993) and Koedijk *et al.* (1996) consider over twenty countries, while Kollmann (1995) considers seven. We choose five countries for the following reason. Among the countries studied in the three above papers, about 80 percent of the countries restricted international financial flows at some point in time, and many did so until the early nineties.[10] These measures can be comparatively mild, like the two-tier exchange rates adopted in Belgium for over 45 years, in the UK during much of the eighties, and for a brief period also in Italy. Other measures have been more obstructive, like licenses or other restrictions (all Southern EU countries including France; also Ireland, Japan), or even

[10] See IFS, various issues, for reviews of exchange restrictions and the like.

full bans on foreign portfolio investment (all Nordic countries, and most NICs and LDCs). Especially the more drastic restrictions make a mockery of tests of a model that assumes perfectly integrated financial markets. Some of the countries with good records in this field have been tracking Germany too closely to provide much additional information, and/or have few world-class firms that effectively attract international investors. In the end, we decided to select the four large economies whose currencies dominate the foreign exchange market and whose securities dominate the world stock market: the USA; the UK, despite a dual exchange rate in the early eighties; Germany; and Japan, despite its exchange controls prior to 1982; and, Switzerland, a mainstream non-ERM economy with unrestricted capital flows.

In the next two sections, our objective is to evaluate the generalized RBC model, that is, the Backus-Smith equation extended so as to allow for noise as well as different time-preference and risk-aversion parameters across countries:

$$S(t)\frac{\Pi_2(t)}{\Pi_1(t)} = e^{\Delta}\left(\frac{C_1(t)}{\Pi_1(t)}\right)^{\eta_1}\left(\frac{C_2(t)}{Pi_2(t)}\right)^{\eta_2}, \tag{7}$$

where Δ is the difference of the time preference parameters, per quarter. The performance of the generalized RBC model is measured relative to is PPP (that is, where $\Delta = \eta_1 = 0 = \eta_2$), and we use data in levels rather than first differences. Our approach is similar to what has become standard in the empirical literature on PPP: rather than requiring that equation (7) hold exactly at any given date, we verify whether the variables identified in the model have an influence on the exchange rate in the long run. If we succeed in doing so, then PPP does not provide the best possible explanation of long-run exchange-rate behavior.

3 Data, estimation procedure, and significance tests

Our data are quarterly consumption spending series, CPI data in the last month of the quarter, and end-of-quarter exchange rate data from IFS for the United States (US), Japan (JP), Germany (DE), the United Kingdom (UK), and Switzerland (CH), over the period 1974:I to 1994:IV. We take the USD as the reference currency (corresponds to currency "1", in the theoretical model) and convert all exchange rates into USD per unit of foreign currency. In what follows, the other country is generally referred to as country $k = \{$DE, JP, US, CH$\}$.

Our regression specification is based on equation (6). We found that no leads were

necessary, and that two quarters amply sufficed at the lag side. Thus, defining $\Delta_l X(t) \equiv X(t-l) - X(t-l-1)$, the equation is

$$\begin{aligned}
\ln\left(S(t)\frac{\Pi_2(t)}{\Pi_1(t)}\right) &= \alpha + \beta_0 t + \beta_k \ln\frac{C_k(t)}{\Pi_k(t)} - \beta_{\mathrm{US}} \ln\frac{C_{\mathrm{US}}(t)}{\Pi_{\mathrm{US}}(t)} \qquad (8)\\
&+ \rho\left\{\ln S(t-1)\frac{\Pi_k(t-1)}{\Pi_{\mathrm{US}}(t-1)} - \alpha + \beta_0(t-1) + \beta_k \ln\frac{C_k(t-1)}{\Pi_k(t-1)} - \beta_{\mathrm{US}} \ln\frac{C_{\mathrm{US}}(t-1)}{\Pi_{\mathrm{US}}(t-1)}\right\}\\
&+ \sum_{L=\mathrm{US},k}\ \sum_{l=1,2} \Delta_l\left(\ln\frac{C_L(t)}{\Pi_L(t)}\right) + \sum_{l=1,2}\Delta_l\left(\ln S(t)\frac{\Pi_k(t)}{\Pi_{\mathrm{US}}(t)}\right) + \epsilon.
\end{aligned}$$

The first line in (8) is the model stated in levels, from equation (7). The second line captures the first-order autocorrelation, ρ, in the deviations from the long-run equilibrium. The third line adds lagged changes in the relevant variables to pick up any remaining predictability in the error.

To estimate (8) we use Seemingly Unrelated Nonlinear Least Squares, first in bilateral estimations (reported in Table 1) where no restrictions are imposed, and then in joint-estimation (Table 2) where we restrict the value of β_{US} (the estimator of η_{US}) to be the same across all the countries. Rather than relying on asymptotic normality, we evaluate the significance of the t-statistics by means of Monte-Carlo experiments. In all of the experiments, price and consumption variables are generated on the basis of estimated VARs, without an error-correction/cointegration term because Johansen-Juselius tests do not reveal any relation among the five price series nor among the five real-consumption series separately. Specifically, the five inflation rates and real-consumption growth rates are first estimated, and then simulated, as mutually correlated ARIMA processes. The simulated exchange-rate data, in contrast, are produced by three alternative "null" data-generating processes, none of which allows a role for real consumption or a time trend. These three "null" data-generating processes are:

(i) ARIMA: the real exchange rates are assumed to be non-stationary, and follow mutually correlated ARIMA (2,1,0) processes (Roll, 1977).

(ii) PPP: the real exchange-rate equation has, in addition to the VAR part, an error-correction term that links the exchange rate to its PPP value. Again, the innovations are correlated across exchange rates.

(iii) Generalized PPP: the procedure is the same as in the previous model, except that now the coefficients δ_{US} and δ_k in the error-correction term for the nominal exchange-rate equation, $\zeta[S(t-1)+\delta_{\mathrm{US}}\Pi_{\mathrm{US}}(t-1)-\delta_k\Pi_k(t-1)]$, are estimated rather than pre-set at unity. This is motivated by the finding of several studies of a cointegration relation between $S(t)$, $\Pi_{\mathrm{US}}(t)$ and $\Pi_k(t-1)]$, with coefficients deviating from unity.

For each data-generating process we simulate 3000 complete 90-quarter, five-country samples (prices, consumptions, and exchange rates), and for each of these samples we estimate equation (8), either bilaterally (in Table 1) or with the constraint that β_{US}, the estimator of η_{US}, be identical across equations (in Table 2). We retrieve the t-statistics for each coefficient, rank them, extract simulated percentile values, and provide values for the $1^{\mathrm{st}}, 5^{\mathrm{th}}, 10^{\mathrm{th}}, 90^{\mathrm{th}}, 95^{\mathrm{th}}$, and 99th percentiles in Tables 1 and 2 under the header "two-sided confidence interval for t-statistics".

4 Empirical results

Table 1 provides the coefficient estimates from the bilateral equations. The header for each country-panel gives the autocorrelation in the deviation from the long-run model, ρ, the R^2 corrected for degrees of freedom,[11] and the Durbin-Watson statistic, DW. For all countries we note that the quarterly deviations from the estimated long-term relation are quite persistent, with ρ being around 0.8. While undoubtedly high, this ρ still implies a half-life of only four to five quarters, which is encouragingly lower than the half-life of PPP deviations estimated to be three to four years in Abuaf and Jorion (1991). The DWs indicate some positive residual autocorrelation. The left-hand side of each country panel shows the estimated coefficients of the long-run model for the exchange rate (β_{US} estimating η_{US}, β_k estimating η_k, and β_0 estimating the difference in the impatience parameters). In the bilateral estimations, six out of eight η-estimates are positive; of the two negative estimates, only one (for German real consumption in the DEM equation) is significant.[12] The average

[11] This includes the part explained by the heavy autocorrelation.

[12] Relative to the standard values, we notice systematically thicker tails in the Monte-Carlo output. Across models, the "Generalized PPP" model tends to generate somewhat wider distributions, while across coefficients the thick tails are especially pronounced for the time-trend coefficient.

risk-aversion coefficient is 1.66.[13] The time trend is significant for the DEM and GBP, and weakly so for the CHF.

The pattern of rejections are similar across the three alternative "null" models. From, for example, Abuaf and Jorion (1991) we know that in a sample of a few decades of monthly data, mean-reversion is hard to detect. What our Monte-Carlo experiments indicate is in line with this: the part of exchange-rate changes picked up by the PPP error-correction model is too small to have a material impact on the structure of the simulated data. The same appears to hold when we allow a relation between $S(t)$, $\Pi_{\mathrm{US}}(t)$ and $\Pi_k(t-1)]$ that is more general than PPP: the generalization seems to have only a marginal impact on the simulated numbers. Thus, irrespective of the details of the null, the generalized RBC model seems to outperform PPP.

The estimates in Table 1 does not restrict the relative risk-aversion coefficient for the US to be the same across equations. The $\chi^2(3)$ statistic on this restriction is 7.429, which would be significant at the 6% level if the distribution were χ^2 indeed. However, simulations reveal that the distribution is substantially wider.[14] From these simulations, the 95% percentile χ^2-test under the null turns out to be 53.84, which is much larger than the asymptotic value of 7.815 in well-behaved samples. Or, summarizing our results in a different way: when using 7.429 as the critical χ^2 value for the hypothesis that the data are generated by a model with equal US coefficient, we observe not 6% incorrect rejections, but 50%. In short, we cannot reject the hypothesis of a common coefficient for the US.

When this common β_{US} is imposed across equations (in Table 2), the results do not change substantially. The average estimate of relative risk aversion across the five estimates is virtually unaffected (1.67). While the estimated coefficient for Japan's real consumption loses significance and the estimated coefficient for German real consumption remains significantly negative, the UK estimate changes sign and becomes positive, and three out of the

[13] To avoid weighting the estimate for the US four times instead of only once, we first compute the average η_{US} over the four separate estimates (which produces $\eta_{\mathrm{US}} = 5.01$) and we then compute an equally-weighted average of the five country estimates. This international average is close to the one in Table 2, where η_{US} is constrained to be equal across all regressions. A simple average, over all eight estimates, would have been 2.92 instead of 1.66.

[14] Given that the null of the χ^2-test is that the generalized RBC model holds with a US coefficient that is identical across equations, the simulated samples are now generated by a generalized RBC model that has as its parameters the estimates from Table 2 (that is, with a common value for η_{US}). From each of the 3000 artificial samples we then estimate the equation without restriction on the US coefficient and we compute a standard χ^2-test for equality of these coefficients across the four equations.

four time-trend coefficients are now significant. In short, we still conclude that both the time-trend and risk-aversion have noticeable effects, and in all cases but one, the estimated risk-aversion parameter has the correct sign. Combined with the conclusions from earlier tests on the presence of cointegration relations, we find that long-run PPP does not provide the best explanation for the level of the exchange rate, and that even a very restrictive model like the RBC one is able to outperform PPP.

5 Conclusion

Much of the literature on exchange rate determination is based on PPP, with PPP being justified on the basis of the consumption opportunity set (frictionless commodity arbitrage). In contrast, the standard micro-economic equilibrium paradigm views relative prices—and, hence, also exchange rates—as determined not just by consumption opportunity sets, but also by marginal utilities. We use regression analysis to test the RBC exchange rate model assuming homothetic, state-independent power utility with constant time-preference parameters; we use levels, an efficient estimation criteria, and Monte-carlo-based significance levels. We find that our model outperforms PPP; that is, real spending and international differences in time preferences appear to have an influence on the real exchange rate.

Appendix: Decomposition of changes in marginal utility

In the derivation of the decomposition, below, we start from the total differential of marginal utility of nominal spending, $\Lambda_k = \Lambda_k(C_k, p_k, X_k, t)$, and then substitute the definition $\Lambda_k = \partial V_k/\partial C_k$ as per equation (4). In the third line we use the definition of time preference, and also invoke the property $\partial V_k/\partial p_{kj} = -c_{kj}\partial V_k/\partial C_k$ (Roy's Identity). We next use the rule for differentiating a product and the definition of ζ_{ks}. Finally, we bring out the percentage changes in the budget and the prices, rearrange, and use the definitions of relative risk aversion η and of total and marginal inflation:

$$
\begin{aligned}
&\frac{d\Lambda_k(t)}{\Lambda_k(t)} \\
&= \frac{1}{\Lambda_k(t)}\left(\frac{\partial\Lambda_k(t)}{\partial t}dt + \frac{\partial\Lambda_k(t)}{\partial C_k}dC_k + \sum_{j=1}^{N}\frac{\partial\Lambda_k(t)}{\partial p_{kj}}dp_{kj} + \sum_{s=1}^{M}\frac{\partial\Lambda_k(t)}{\partial X_{ks}}dX_{ks}\right) \\
&= \frac{\partial^2 V_k/(\partial C_k\partial t)}{\partial V_k/\partial C_k}dt + \frac{1}{\partial V_k/\partial C_k}\left(\frac{\partial^2 V_k}{\partial C_k^2}dC_k + \sum_{j=1}^{N}\frac{\partial^2 V_k}{\partial p_{kj}\partial C_k}dp_{kj} + \sum_{s=1}^{M}\frac{\partial^2 V_k}{\partial C_k\partial X_{ks}}dX_{ks}\right) \\
&= -\delta_k(t)dt + \frac{1}{\partial V_k/\partial C_k}\left(\frac{\partial^2 V_k}{\partial C_k^2}dC_k - \sum_{j=1}^{N}\frac{\partial\frac{\partial V_k}{\partial C_k}c_{kj}}{\partial C_k}dp_{kj} + \sum_{s=1}^{M}X_{ks}\frac{\partial^2 V_k}{\partial C_k\partial X_{ks}}\frac{dX_{ks}}{X_{ks}}\right) \\
&= -\delta_k(t)dt + \frac{1}{\partial V_k/\partial C_k}\left(\frac{\partial^2 V_k}{\partial C_k^2}dC_k - \sum_{j=1}^{N}\left[\frac{\partial^2 V_k}{\partial C_k^2}c_{kj} + \frac{\partial V_k}{\partial C_k}\frac{\partial c_{kj}}{\partial C_k}\right)dp_{kj}\right] + \sum_{s=1}^{M}\zeta_{ks}\frac{dX_{ks}}{X_{ks}} \\
&= -\delta_k(t)dt - \frac{-C_k\partial^2 V_k/\partial C_k^2}{\partial V_k/\partial C_k}\left(\frac{dC_k}{C_k} - \sum_{j=1}^{N}\frac{c_{kj}p_{kj}}{C_k}\frac{dp_{kj}}{p_{kj}}\right) - \sum_{j=1}^{N}\frac{\partial c_{kj}}{\partial C_k}p_{kj}\frac{dp_{kj}}{p_{kj}} + \sum_{s=1}^{M}\zeta_{ks}\frac{dX_{ks}}{X_{ks}} \\
&= -\delta_k(t)dt - \eta_k(t)\left(\frac{dC_k}{C_k} - \frac{d\Pi_k}{\Pi_k}\right) - \frac{d\pi_k}{\pi_k} + \sum_{s=1}^{M}\zeta_{ks}\frac{dX_{ks}}{X_{ks}} \qquad \text{(A1)}
\end{aligned}
$$

Substitution of (A1) into (1) then gives (5).

Table 1: Regression with β_{US} *not* restricted to be equal across regressions

coeff	• estimate • std. error • t-stat	α	Two-sided confidence intervals for t-statistic under three alternative null hypothesis								
			ARIMA			PPP			Generalized PPP		
Deutsche Mark ($\rho = 0.809$, $R^2 = 0.913$, DW $= 1.766$)											
β_k	-1.361	10%	-1.4238	1.3678	*	-1.4543	1.4824	*	-1.5026	1.5294	*
	0.350	5%	-1.8071	1.782	*	-1.861	1.9053	*	-1.872	2.0048	*
	-3.889	1%	-2.6145	2.6688	*	-2.6999	2.6906	*	-2.631	3.0556	*
β_{US}	4.909	10%	-1.5457	1.5013	*	-1.393	1.3616	*	-1.3605	1.3356	*
	1.220	5%	-1.9639	1.9697	*	-1.7605	1.743	*	-1.7753	1.7771	*
	4.025	1%	-2.7592	2.8371	*	-2.5636	2.5052	*	-2.6348	2.8066	*
β_0 $(E-3)$	7.333	10%	-1.5999	1.6167	*	-1.6675	1.4276	*	-1.704	1.4903	*
	1.638	5%	-2.1142	2.1225	*	-2.1488	1.8659	*	-2.1656	1.9418	*
	4.476	1%	-2.8536	2.9706	*	-3.0877	2.5962	*	-2.9974	2.8354	*
British Pound ($\rho = 0.749$, $R^2 = 0.878$, DW $= 1.25$)											
β_k	-0.449	10%	-1.4864	1.408		-1.3621	1.4523		-1.2772	1.294	
	0.848	5%	-1.9068	1.8204		-1.7429	1.8496		-1.649	1.6888	
	-0.529	1%	-2.9316	2.8013		-2.7553	2.754		-2.9051	2.7175	
β_{US}	2.255	10%	-1.4444	1.4277	*	-1.3802	1.3952	*	-1.3431	1.2993	*
	1.157	5%	-1.8311	1.8053	*	-1.7866	1.7736	*	-1.7744	1.7239	*
	1.949	1%	-2.8189	2.8311		-2.7591	2.6863		-3.2453	3.0464	
β_0 $(E-3)$	3.861	10%	-1.6305	1.6711	*	-1.6244	1.6426	*	-2.0088	1.9769	*
	1.219	5%	-2.0951	2.1556	*	-2.1509	2.1161	*	-2.543	2.538	*
	3.167	1%	-2.9844	2.9877	*	-3.1726	2.9394	*	-3.574	3.6704	
Japanese Yen ($\rho = 0.870$, $R^2 = 0.947$, DW $= 1.824$)											
β_k	3.256	10%	-1.7155	1.6789		-1.5112	1.4351	*	-1.3952	1.3958	*
	2.064	5%	-2.1117	2.1973		-1.9593	1.8868		-1.9177	1.9399	
	1.577	1%	-2.9803	3.5422		-2.9972	2.9878		-4.2616	4.4349	
β_{US}	8.561	10%	-1.5926	1.6029	*	-1.4537	1.4133	*	-1.2798	1.2714	*
	2.792	5%	-2.0162	2.047	*	-1.8997	1.8466	*	-1.7776	1.8421	*
	3.066	1%	-3.0776	2.9235	*	-2.8317	2.7139	*	-3.016	3.411	
β_0 $(E-3)$	2.717	10%	-1.7719	1.7438		-1.7478	1.6903		-1.7437	1.8014	
	2.490	5%	-2.2415	2.2037		-2.3547	2.2102		-2.2049	2.2714	
	1.091	1%	-3.178	3.1823		-3.2268	3.2893		-3.3929	3.3628	
Swiss Franc ($\rho = 0.787$, $R^2 = 0.883$, DW $= 1.664$)											
β_k	1.851	10%	-1.5963	1.5943	*	-1.4874	1.4531	*	-1.4829	1.5268	*
	0.842	5%	-2.032	2.0554	*	-1.8319	1.855	*	-1.9734	1.9875	*
	2.199	1%	-2.8545	3.057		-2.5657	2.8291		-3.2503	3.2773	
β_{US}	4.304	10%	-1.5979	1.6075	*	-1.4491	1.4611	*	-1.5009	1.3872	*
	1.219	5%	-2.0257	2.0258	*	-1.8264	1.8425	*	-1.9731	1.8311	*
	3.532	1%	-2.837	2.9465	*	-2.5078	2.6509	*	-3.286	2.9526	*
β_0 $(E-3)$	3.171	10%	-1.7438	1.7632	*	-1.6633	1.5952	*	-2.063	1.6452	*
	1.616	5%	-2.2912	2.2043		-2.14	2.0368		-2.6023	2.2288	
	1.963	1%	-3.0705	3.1256		-3.0278	3.1257		-3.6843	3.1924	

Please see next page for notes to table ...

Notes for Table 1

The table reports results from estimating equation (8) using Seemingly Unrelated Nonlinear Least Squares without any cross-equation constraints on β_{US}. The reference country is the US, the foreign countries (k) are Germany, UK, Japan, and Switzerland. The log level of each real exchange rate, $\ln[S(t)\Pi_k(t)/\Pi_{\mathrm{US}}(t)]$, is regressed on a constant, a time trend (whose coefficient, β_0, estimates the difference in impatience), and the log of domestic and foreign real consumptions, $\ln[C_L(t)/\Pi_L(t)]$, whose coefficients β_L estimates relative risk aversion. According to the model, $\alpha = \ln[\theta_k(1-\eta_k)/(1-\eta_{\mathrm{US}})](>0)$, $\beta_0 = \delta_{\mathrm{US}} - \delta_k(\gtrless 0), \beta_{\mathrm{US}} = \eta_{\mathrm{US}}(>0)$, and $\beta_k = \eta_k(>0)$. Also included in the regression is the beginning-of-period deviation between the exchange rate and its theoretical value), and lagged changes in the regressors and regressand. Data are quarterly, 1974:04 to 1996:02 (87 data points after correcting for lags), from IFS. Exchange rates are end-of-quarter. Consumption C_k is private consumption, Π_k is the end-of-quarter CPI. The header for each country panel shows the autocorrelation in the deviation from the long-run model, ρ, the R^2 corrected for degrees of freedom, and the Durbin-Watson statistic, DW. The body of each country-panel shows the estimated coefficients of the long-run model for the exchange rate, as well as confidence intervals for the t-statistics, two-sided, for $2\alpha = 10$, 5, and 1 percent. These confidence intervals are obtained from Monte-Carlo simulations, as explained in the text. Asterisks beside an interval indicate that the observed t-statistic is outside the interval—that is, the assumed data-generating process is rejected.

Table 2: Regression with β_{US} restricted to be equal across all countries

coeff	• estimate • std. error • t-stat	α	Two-sided confidence intervals for t-statistic under three alternative null hypothesis								
			ARIMA			PPP			Generalized PPP		
Deutsche Mark ($\rho = 0.827$, $R^2 = 0.913$, DW = 1.798)											
β_k	-1.401	10%	-1.45	1.396	*	-1.434	1.464	*	-1.509	1.544	*
	0.381	5%	-1.837	1.838	*	-1.982	2.187	*	-2.002	1.977	*
	-3.679	1%	-2.604	2.594	*	-4.001	5.442		-2.846	2.794	*
$\beta_0(\text{E}-3)$	7.394	10%	-1.929	1.879	*	-2.167	2.099	*	-2.096	1.843	*
	1.620	5%	-2.46	2.359	*	-2.89	2.745	*	-2.574	2.378	*
	4.565	1%	-3.197	3.159	*	-4.273	4.22	*	-3.524	3.3	*
British Pound($\rho = 0.842$, $R^2 = 0.878$, DW = 1.828)											
β_k	1.17	10%	-1.446	1.509		-1.934	2.148		-1.3	1.302	
	1.055	5%	-1.855	1.92		-2.567	2.737		-1.761	1.727	
	1.109	1%	-2.53	2.819		-3.848	4.055		-2.761	2.766	
$\beta_0(\text{E}-3)$	4.113	10%	-1.793	1.774	*	-2.796	2.982	*	-2.098	2.035	*
	1.202	5%	-2.294	2.325	*	-3.598	3.79		-2.697	2.624	*
	3.420	1%	-3.486	3.115	*	-5.674	5.884		-3.853	3.701	
Japanese Yen ($\rho = 0.831$, $R^2 = 0.946$, DW = 1.756)											
β_k	1.647	10%	-1.604	1.622		-1.628	1.703		-1.312	1.343	
	1.723	5%	-1.986	2.081		-2.233	2.283		-1.734	1.811	
	0.956	1%	-3.105	2.974		-3.52	3.87		-1.312	3.496	
$\beta_0(\text{E}-3)$	2.556	10%	-1.953	1.901		-1.898	2.009		-1.876	1.851	
	2.469	5%	-2.406	2.394		-2.424	2.681		-2.456	2.419	
	1.035	1%	-3.365	3.317		-3.736	3.95		-3.479	3.488	
Swiss Franc ($\rho = 0.827$, $R^2 = 0.$ 0.883, DW = 1.705)											
β_k	1.328	10%	-1.541	1.611		-2.555	2.581		-1.445	1.417	
	1.037	%5	-1.982	2.065		-3.348	3.509		-1.899	1.857	
	1.280	1%	-2.797	2.799		-5.022	5.539		-2.82	2.709	
$\beta_0(\text{E}-3)$	4.361	10%	-2.011	1.915	*	-3.252	3.541		-2.228	1.829	*
	1.452	5%	-2.511	2.434	*	-4.301	4.473		-2.766	2.385	*
	3.002	1%	-3.444	3.396		-6.093	6.383		-3.931	3.466	
United States (common to all equations)											
β_{US}	5.608	10%	-1.942	1.994	*	-2.215	2.228	*	-1.898	1.981	*
	1.292	5%	-2.469	2.424	*	-2.922	2.855	*	-2.362	2.419	*
	4.338	1%	-3.424	3.365	*	-4.239	4.323	*	-3.374	3.346	*

Please see next page for notes to table ...

Notes for Table 2

The table reports results from estimating equation (8) using Seemingly Unrelated Nonlinear Least Squares *with* the restriction that $\beta_{\text{US}} = \eta_{\text{US}}(> 0)$ must be the same across all countries. The reference country is the US, the foreign countries (k) are Germany, UK, Japan, and Switzerland. The log level of each real exchange rate, $\ln[S(t)\Pi_k(t)/\Pi_{\text{US}}(t)]$, is regressed on a constant, a time trend (whose coefficient, β_0, estimates the difference in impatience), and the log of domestic and foreign real consumptions, $\ln[C_L(t)/\Pi_L(t)]$, whose coefficients β_L estimates relative risk aversion. According to the model, $\alpha = \ln[\theta_k(1-\eta_k)/(1-\eta_{\text{US}})](> 0)$, $\beta_0 = \delta_{\text{US}} - \delta_k(\gtrless 0)$, and $\beta_k = \eta_k(> 0)$. Also included in the regression is the beginning-of-period deviation between the exchange rate and its theoretical value, and lagged changes in the regressors and regressand. Data are quarterly, 1974:04 to 1996:02 (87 data points after correcting for lags), from IFS. Exchange rates are end-of-quarter. Consumption C_k is private consumption, Π_k is the end-of-quarter CPI. The header for each country panel shows the autocorrelation in the deviation from the long-run model, ρ, the R^2 corrected for degrees of freedom, and the Durbin-Watson statistic, DW. The body of each country-panel shows the estimated coefficients of the long-run model for the exchange rate, as well as confidence intervals for the t-statistics, two-sided, for $2\alpha = 10$, 5, and 1 percent. These confidence intervals are obtained from Monte-Carlo simulations, as explained in the text. Asterisks beside an interval indicate that the observed t-statistic is outside the interval—that is, the assumed data-generating process is rejected.

References

Abuaf, Niso and Philippe Jorion (1990): Purchasing Power Parity in the Long Run, *Journal of Finance* 45.1, 157-174.

Backus, D., S. Foresi and C. Telmer (1996): Affine Models of Currency Pricing, NYU Working paper.

Backus, D. and G. Smith (1993): Consumption and Real Exchange Rates in Dynamic Economies with Non-traded Goods, *Journal of International Economics* 35, 297-316.

Bakshi, G. and Z. Chen (1997): Equilibrium Valuation of Foreign Exchange Claims, *Journal of Finance* 52.2, 799-826.

Basak, S. and M. Gallmeyer (1999): Currency Prices, the Nominal Exchange Rate, and Security prices in a Two Country Dynamic Monetary Equilibrium, *Mathematical Finance* 9, 1-30.

Breeden, D. (1979): An Intertemporal Asset Pricing Model with Stochastic Consumption and Investment Opportunities, *Journal of Financial Economics* 7, 265-296.

Devereux, M. (1997): Real Exchange Rates and Macroeconomics: Evidence and Theory,*Canadian Journal of Economics* 30, 773-808.

Duffie, D. (1992): *Dynamic Asset Pricing Theory*, Princeton University Press, Princeton NJ.

Dumas, B. (1992): Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World, *The Review of Financial Studies* 5, 153-180.

Edison, Hali, Joseph Gagnon, and William Melick (1997): Understanding the Empirical Literature on Purchasing Power Parity: The Post-Bretton Woods Era, *Journal of International Money and Finance* 16.1, 1-17.

Engel, C. (1993): Real exchange rates and relative prices: An empirical investigation, *Journal of Monetary Economics* 32.1, 35-50.

Engel, C. and J. Rogers (1995): How wide is the border?, *American Economic Review* 86.5, 1112-1125.

Engel, C., M. Hendrickson and J. Rogers (1996): Intra-National, Intra-Continental, and Intra-Planetary PPP, U. of Washington Working paper.

Frankel, J. A. and A. K. Rose, (1996): A Panel Project on Purchasing Power Parity: Mean Reversion Within and Between Countries, *Journal of International Economics* 40, 209-224.

Froot, K. A., M. Kim and K. Rogoff (1995): The Law of One Price over 700 Years, NBER Working paper No. 5132.

Froot, K. A. and K. Rogoff (1995): Perspectives on PPP and Long-Run Real Exchange Rates, in *Handbook of international Economics* 3, G. Grossman and K. Rogoff, eds. Amsterdam, Elsevier-North Holland.

Hollifield, B. and R. Uppal (1997): An Examination of Uncovered Interest Rate Parity in Segmented International Commodity Markets, *Journal of Finance*, forthcoming.

Horvath, M. and M. Watson (1995): Testing for Cointegration When Some of the Cointegrating Vectors are Known, *Econometric Theory* 11, 984-1014.

Johansen, S. and K. Juselius (1992): Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and the UIP for UK, *Journal of Econometrics* 53, 211-244.

Koedijk, C., F. Nissen, P. Schotman, C. Wolff (1996): PPP and Real Expenditure, Working paper, Maastricht University.

Kollmann, R. (1995): Consumption, Real Exchange Rates and the Structure of International Asset Markets, *Journal of International Money and Finance* 14.2, 191-211.

Lane, Philip R. (1999): The New Open Economy Macroeconomics: a Survey, CEPR#2115.

Lothian, J. (1997): Multi-country Evidence on the Behavior of Purchasing Power Parity under the Current Float, *Journal of International Money and Finance* 16.1, 19-35.

Lothian, J. and M. Taylor (1996): Real Exchange Rate Behavior: The Recent float from the Perspective of the Past Two Centuries, *Journal of Political Economy* 104 , 488-509.

Lucas, R. (1982): Interest Rates and Currency Prices in a Two-Country World, *Journal of Monetary Economics* 10, 335-359.

Nessén, M. (1994): Common Trends in Prices and Exchange Rates-Tests of Long-Run Purchasing Power Parity, Essay I in Essays on Exchange Rates and International Finance, Unpublished Doctoral Dissertation, Department of Finance, Stockholm School of Economics.

Nielsen, L. and Saá-Requejo (1993): Exchange Rate and the Term Structure of Risk Premia in foreign Exchange Markets, University of Chicago Working paper.

O'Connell, P. (1998): The Overvaluation of Purchasing Power Parity, *Journal of International Economics* 44.1, 1-19.

Obstfeld, Maurice and Kenneth Rogoff (2000): New Directions for Stochastic Open Economy Models, *Journal of International Economics* 50, 117-154.

Phillips, P.C.B. and M. Loretan (1991): Estimating Long-run Economic Equilibria, *Review of Economic Studies* 58, 407-436.

Rogers, John and Michael Jenkins (1995): Haircuts or hysteresis? Sources of movements in real exchange rates, *Journal of International Economics* 38, 339-360.

Rose, A. (1996/7): International Finance and Macroeconomics, NBER Reporter Winter, 1-7.

Sercu, P., R. Uppal and C. Van Hulle (1995): The Exchange Rate in the Presence of Transactions Costs: Implications for Tests of Purchasing Power Parity, *Journal of Finance* 50, 1309-1319.

Sercu, P. and R. Uppal (2000): *Exchange Rate Volatility, Trade, and Capital Flows under Alternative Exchange Rate Regimes*, Cambridge University Press, Cambridge UK.

Stockman, A. (1980): A Theory of Exchange Rate Determination, *Journal of Political Economy* 88, 673-698.

Stulz, R. (1987): An Equilibrium Model of Exchange Rate Determination and Asset Pricing with Nontraded Goods and Imperfect Information, *Journal of Political Economy* 95, 1024-1040.

Taylor, A., (1995): International Capital Mobility in History: Purchasing Power Parity in the Long Run, NBER Working paper No. 5675.

Wei, J. and D. Parsley, (1995): Purchasing Power Disparity During the Floating Rate Period: Exchange Rate Volatility, Trade Barriers and Other Culprits, NBER Working paper No. 5032.