1. Introduction

Estimates of 'equilibrium' values of real exchange rates often play an important role in macroeconomic policy debates, both under fixed and flexible exchange rate regimes. Although PPP based estimates are frequently used in this context, there remain serious doubts about their theoretical and empirical validity, particularly over a medium term time horizon. A number of alternatives to PPP have been proposed, the most widely used of which are concepts related to Williamson's Fundamental Equilibrium Exchange Rate (hereafter FEER).[1] Although there exists an extensive literature on testing PPP, there has been relatively little research on testing alternatives.

FEERs can be calculated in a number of ways, but the most widely used approach is based on a partial equilibrium analysis, which inputs exogenous estimates of trend output and structural capital flows into a model of aggregate trade. (For an extensive review of this and other approaches, see MacDonald and Stein, 1999). Estimates using this method have been influential in a number of policy debates (e.g. Williamson, 1991). However, FEER estimates usually relate to a specific point in time, so the approach has never been systematically tested in a time series context.

In this paper we present a test of FEER estimates for six major industrialised countries based on cointegration analysis. If PPP does not hold over the medium term, then real exchange rates over the last 20/30 years rate might well be non-stationary, and this is the case for the data we use. A natural question to ask of FEERs is whether they can account for this non-stationary behaviour i.e. whether FEER estimates and real exchange rates cointegrate. We examine this for each country both individually, and collectively using panel unit root tests.

Section 2 outlines the partial equilibrium approach to estimating FEERs, and examines the implications for the time series properties of FEERs and the real exchange rate. Section 3 estimates time series for FEERs for six economies from the mid-70s. Section 4 examines whether the FEER and the real exchange rate cointegrate. Section 5 concludes.

[1] Williamson, 1983.

2. Partial Equilibrium Calculations of FEERs

In this section we outline the FEER concept and our implementation of it. Our aim is not to justify or defend the approach: indeed we note a number of criticisms and weaknesses below. In our view the widespread use of FEER estimates in policy debates is sufficient justification for attempting to test FEERs. What is important is that, in deriving historic series for FEERs, we follow and respect the literature and methodology involved in making FEER calculations.

There are two defining characteristics of the FEER approach. The first is the proposition that most trade in developed economies involves differentiated goods sold in imperfectly competitive markets. As a result individual firms, and economies as a whole, face a downward sloping demand curve for their products. This is in direct contrast to PPP, where perfect arbitrage is assumed to force prices, and therefore real exchange rates, back to some constant level. The second feature of FEERs is that they apply to a medium term horizon. FEERs abstract from short term disequilibria caused, for example, by expectations errors or nominal inertia, with the result that monetary policy cannot influence the FEER. However the FEER is not a long run equilibrium concept. Estimates of medium term, or ‘structural’, capital flows play a crucial role in FEER estimates so that asset stocks are changing. In this sense the FEER relates to flow, not stock, equilibrium.

The basic FEER approach can be illustrated using the following aggregate equations for goods market equilibrium and the balance of payments:

$$Y(K(i), N, ..) = DD(...) + NT(WT, e, Y, rcp, ..) \qquad (1)$$

$$NT(WT, e, Y, rcp, ..) - iD + (i_W + \Delta e/e) Ae = \Delta(Ae) - \Delta D \qquad (2)$$

where Y is real output, DD domestic demand (private and public consumption and investment), K the capital stock, N labour supply, NT net trade, WT world trade, e the real exchange rate (measured as domestic currency per unit of foreign), rcp are real commodity prices, i is the domestic real interest rate, D is the real value of debt issued domestically and denominated in domestic currency held by overseas residents, A is the real value of overseas debt denominated in overseas currency held by domestic residents, and i_W is the overseas real interest rate.

Equation (1) equates output supply to the demand for domestic output so that there is ‘internal balance’. This assumption clearly abstracts from Keynesian or other types of short term disequilibria. If we make the conventional assumption that the model is superneutral, then internal balance implies that we can abstract from monetary policy and the determination of inflation in evaluating the equilibrium real exchange rate. (It also implies, of course, that the FEER is a real rather than a nominal exchange rate).

As equation (2) shows, asset stocks are not assumed to be constant. This is based on the view that asset adjustment can in practice be drawn out over many decades (for example, as a result of intertemporal consumption smoothing) and so an equilibrium exchange rate concept which applies only to a long run stock equilibrium would be of limited practical value.

The FEER approach is quite eclectic, and can encompass quite different models of, say, consumption or investment. The current account surplus or deficit will reflect both private sector net saving decisions and the fiscal policy stance. The latter has led some to describe the FEER as a normative concept (and occasionally re-christened it as the 'Desired EER'). Thus, although the FEER is independent of monetary policy, it will be influenced by fiscal policy decisions.[2]

One important aspect is the role of imperfect competition which is reflected in the assumption that net trade is a function of the real exchange rate. For the same reason, the appropriate definition of the real exchange rate is one involving relative output or traded goods prices rather than consumer prices. This distinction is important when countries produce different goods. In Obstfeld and Rogoff (1995), for example, PPP holds for relative consumer prices but not relative producer prices.[3] Within this framework, the real exchange rate will only be constant (and in this sense PPP will only hold), if trade elasticities are very large. PPP becomes a special case of the model. In the general case, movements in any one of the variables of the model will generate movements in the real exchange rate.

The stationarity of the real exchange rate has been used extensively to assess the validity of PPP. Within the context of this model, we would only expect the real exchange rate to be stationary if all exogenous influences on the model are stationary.

[2] The independence of real interest rates from monetary policy does not imply that real interest rates must always equal world real rates (+/- a risk premium), as real exchange rates may be changing over time, because of asset accumulation for example: see Driver and Wren-Lewis, 1998.

[3] This occurs as all consumers are identical across countries, there is no home bias in consumption, there is no pricing-to-market and the law of one price holds for each individual good. If any of these assumptions are relaxed, any non-stationarity in relative output prices would also manifest itself in relative consumer prices.

If the observed real exchange rate is non-stationary, we would nevertheless expect the model's prediction errors for the real exchange rate (i.e. the difference between the actual real rate and the FEER) to be stationary. This in effect assumes that the short term influences ignored by the FEER model (e.g. Keynesian disequilibria and expectational errors) are stationary.[4]

There are three main approaches to estimating FEERs in the literature. The first uses a complete macroeconomic model, where for example the DD(..) function is fully specified, and generates the FEER as a solution to that model (see Williamson, 1994). This general equilibrium approach has its advantages and disadvantages. On the one hand the estimated FEER will reflect a complete set of endogenous feedbacks. On the other, estimating a complete model from scratch is quite costly, while adapting an existing model may lead to problems in interpretation.

A second method directly estimates a relationship for the real exchange rate, including as explanatory variables determinants of the variables in (1) and (2) (see Stein, 1994, for example). Reflecting the fact that in general the real exchange rate is found to be I(1), this type of analysis uses cointegration techniques. This type of approach has the advantages and disadvantages of any reduced form methodology.

The third and most widely used method to calculate FEERs takes a partial equilibrium approach. This involves taking off-model estimates of trend output Y, world trade WT, commodity prices and the medium term asset accumulation, and inverting (2) to solve for the real exchange rate.[5] The partial equilibrium approach also typically takes the value of asset stocks (A and D) as given, so that debt interest flows are endogenous only to the extent of revaluation effects caused by movements in the exchange rate. (It is as if the FEER represents the real exchange rate that would occur if we immediately moved to internal balance, with no transition path for asset stocks).[6]

The partial equilibrium approach has the advantage that only a part of the macroeconomy needs to be estimated – the net trade function NT(…). The major

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[4] If the actual real rate is stationary, this does not immediately imply that the FEER approach is incorrect, because it could be, for example, that all the exogenous variable inputs into the FEER are also stationary. However, this case does not turn out to be relevant for the real exchange rate series examined in this paper.

[5] The FEER is often described as the real exchange rate that equates the current account to medium term or structural capital flows. However, as (1) and (2) show, this is equivalent to evaluating the level of domestic demand consistent with assumptions about trend output and asset accumulation, and calculating the FEER as the real exchange rate that equates the demand and supply of domestic output.

[6] The sensitivity of FEER estimates to this assumption has been explored briefly in Wren-Lewis, 1992, and in more detail in Artis and Taylor, 1995.

disadvantage is that inconsistencies may arise between off-model assumptions and the solution for the real exchange rate. The greater the feedback from the real exchange rate to trend output or saving and investment decisions, the greater this danger is. These concerns are discussed and analysed in Driver and Wren-Lewis, 1999a.

In this paper we will examine the partial equilibrium approach to estimating FEERs. We therefore estimate an aggregate trade model for six major industrialised economies, adopting an approach similar to that used in the literature. The NT( ) function is disaggregated into four components: into prices and quantities, and exports and imports. Some studies involve an extra element of disaggregation, distinguishing between trade in goods and trade in services, but we use aggregate trade in goods and services equations. We also adopt the standard specification in the literature:

$$NT(\ldots) = X(WT,e/rpx) * rpx(e,rcp_X) - M(Y,e/rpm) * rpm(e,rcp_M) \qquad (3)$$

where X is the volume of exports, WT is world trade volume, e is the real exchange rate (defined as the nominal rate multiplied by ‘world’ export prices divided by domestic output prices), rpx are real export prices (export prices in domestic currency divided by output prices), rcp real commodity prices, M is the volume of imports, Y is real domestic output, and rpm are real import prices. (Precise data definitions are given in Appendix A).

The trade volume equations embody the traditional ‘demand curve’ approach (see Goldstein and Kahn, 1985, for example) that remains the standard way of modelling real trade flows. Trade prices depend on commodity prices, domestic output prices and world export prices. Detailed specification for each country are discussed in the next section.

The partial equilibrium approach takes as inputs off-model estimates of trend output and structural capital flows. As these studies typically derive estimates for FEERs at one particular point in time[7], inputs can be derived in a fairly eclectic manner. In our case, however, we require time series for these inputs to produce time series for FEERs, so we are forced to take a more specific approach. In the case of trend output, we derive estimates using the Hodrick-Prescott filter. This technique for deriving trend output is widely used: see Giorno *et al.* (1995) for example, and often

[7] Barrell and Wren-Lewis, 1989, is an exception.

provides an input into specific FEER studies. We use the same method to derive trend commodity prices (see next section).

We adopted a different technique for structural capital flows for two reasons. First, the time series characteristics of current accounts are such that different parameterisations of the filter could give quite different results. Second, as Driver and Wren-Lewis, 1999a, show estimates of FEERs are much more sensitive to current account assumptions than inputs on trend output. Instead we adopted an econometric approach which focuses on two determinants of structural capital flows that have been emphasised in the literature. (For a similar approach, see Isard and Farquee, 1998). First, differences in age structures between nations have been cited as a key factor behind persistent US current account deficits and Japanese surpluses. Second, the fiscal stance of the government is likely to be a key determinant of national net saving. Our panel estimates of structural capital flows are discussed in more detail below. While this procedure goes beyond normal FEER studies, it stays within the spirit of the partial equilibrium approach, as our model of structural capital flows is independent of the solution for the real exchange rate.[8] It is also a key component of the FEER style macroeconomic balance approach to exchange rate assessment adopted by the IMF, see Isard and Faruqee, 1998.

Finally it is important to remember the current account is not just made up of net trade flows. We follow the literature in keeping IPD flows exogenous, except that we allow for exchange rate revaluations and smooth series using the HP filter.[9] We also smooth the final component of the current account, net transfers, using an HP filter.

[8] The validity of this recursive structure is discussed in Driver and Wren-Lewis, 1998.

[9] At first sight it may seem odd to estimate a model of trend asset flows, but treat assets stocks as fixed. In forecasting FEERs over the future, Driver and Wren-Lewis, 1998, do integrate stock and flow assumptions. Unfortunately this is not an option in deriving historic FEERs, as there is no way of choosing a base date from which to cumulate structural capital flows to obtain asset stocks.

3. Estimating FEERs

In parameterising equation (3) we wanted to keep the estimation process as simple as possible, in part to avoid any suspicion that equation selection had been influenced by the tests of FEERs that follow. As the FEER describes a medium term equilibrium, our concern is not the short run dynamics of these equations, but their longer term properties. For the trade equations, as the tests show that the variables are I(1), static OLS (i.e. Engle and Granger, 1987) seems an appropriate and simple estimation procedure. For the trade volume equations we also look at the long run solutions from dynamic error-correction specifications, and at equation stability. Tables 1 and 2 summarise the selected long run parameters.[10]

Table 1 Trade Volume Equations

Country	Exports			Imports		
	Activity	Comp	Trend	Activity	Comp	Trend
France	1.0*	0.64	>90 1.5%	1.65	-0.29	-
Germany	1.0*	0.52	-	1.68	-0.74*	-
Italy	1.0*	0.53	-	1.51	-0.35	-
Canada	1.0*	0.33	-	1.39	-0.40	>85 3.4%
US	1.0*	1.07	-	2.00	-0.50	-
Japan	1.0*	0.95	<84 2.5%	1.00	-0.29	-
UK	1.0*	1.13	-	2.07	-0.20*	-

All equations estimated from 1973Q1 to 1997Q4. A * indicates an imposed elasticity, while >90 1.5% indicates a split time trend which adds 1.5% p.a. from 1990. Both German equations included a shift dummy from 1991Q1.

In the case of export volumes, for all countries we impose a trade share form on the equation by restricting the world trade elasticity to unity before freely estimating the competitiveness elasticity. The resulting competitiveness elasticities range in size from 0.33 to 1.13, which although small is typical of the literature. With import volumes the activity measure is GDP, so a share specification is not appropriate, and estimated activity elasticities are above unity, as we would expect. In two cases the import competitiveness elasticities had to be imposed: for Germany, where the imposed coefficient comes from a dynamic error correction model (the freely estimated static OLS value is slightly positive); and for the UK, where an insignificant,

[10] Driver and Wren-Lewis, 1998, also examine Johansen estimates, and consider in detail whether these equations cointegrate. Our concern here, however, is to test the properties of FEER estimates themselves, rather than to test each aspect of the models used to derive FEERs.

small positive parameter is replaced by an elasticity estimated in Driver and Wren-Lewis, 1998.

A few of the volume equations required split trends in order to obtain sensible results. Chow, 1960, tests indicate a structural break for French export volumes at the end of the 1980s, with a trend fall in export share thereafter. Japanese exports grew rapidly at the beginning of the estimation period, but Chow tests indicate that this period of rapid growth came to an end around 1984. We also require a positive trend in Canadian imports after 1985. A 'unification' dummy is included for both German trade volume equations.

Taken as a whole, these elasticities are close to the Marshall Lerner condition for France, Canada and Italy. However, once the effect of the exchange rate in revaluing external assets is taken into account, the Marshall Lerner condition holds for all countries, although by only a small margin for Italy. Driver and Wren-Lewis, 1998, obtained a similar result.

Table 2 Trade Price Equations

Country	Exports			Imports		
	World	Commod	Trend	World	Commod	Trend
France	0.27	0.10	-	0.56	0.09	.00005
Germany	0.19	0.10	-	0.49	0.04	-
Italy	0.27	0.11	-.0010	0.57	0.20	-
Canada	0.49	0.15	-.0018	0.42	0.06	-.0038
US	0.28	0.08	-.0040	0.68	0.18*	.0003
Japan	0.42	0.01	-.0043	0.75	0.17	-
UK	0.30	0.14	-	0.42	0.20	-

All equations estimated from 1973Q1 to 1997Q4 except US exports prices which start from 1980Q1, Japan imports prices from 1981Q1.A * indicates an imposed elasticity

In the case of trade prices, we include a simple trend in a number of the specifications. Trend divergences between total output prices and traded goods prices could occur for a number of reasons, such as different rates of growth in productivity in manufacturing and services. Commodity prices for imports and exports are defined as a weighted average of five IFS commodity prices series (see Appendix A), where country specific weights based on the relevant shares are used. Only in the case of US imports is it necessary to impose the coefficient on commodity prices in the trade price equation, as the freely estimated parameter is incorrectly signed.[11]

[11] The commodity composition of trade share could have been used to impose the coefficient on commodity prices, although it is unclear which year to choose. The parameters on commodity

Our estimates of trend output for each country are derived by applying a Hodrick-Prescott filter (see Hodrick and Prescott, 1997). The Hodrick-Prescott filter approach fits a trend through all observations, that minimises a weighted average of the gap between output, Y_t and trend output, Y_t^* at any point in time, and the rate of change in trend output at that point in time. More precisely, the trend output Y_t^* for t = 1,2.....T is estimated to minimise:

$$\sum_{t=1}^{T}(\ln Y_t - \ln Y_t^*)^2 + l \sum_{t=2}^{T-1}\left[(\ln Y_{t+1}^* - \ln Y_t^*) - (\ln Y_t^* - \ln Y_{t-1}^*)\right]^2 \qquad (4)$$

where the parameter l controls the smoothness of the resulting trend line. A l close to zero will produce a trend that follows output more closely. Conversely, as l tends to infinity the Hodrick Prescott trend approaches a linear time trend. The trend is computed by applying the Hodrick-Prescott technique with the smoothing factor $l = 1600$ to real GDP data[12]. Giorno *et al.*, 1995, argue that this type of filter is subject to an "end point" problem. This partly reflects the fitting of a trend line symmetrically through the data: if the beginning and the end of the data set do not reflect similar points in the cycle then the trend will be pulled upwards or downwards towards the path of actual output for the first few and the last few observations. Following Giorno *et al.*, 1995, in order to reduce the end-point problem and to give more stability to estimates the series are constructed by extending the period of estimation until 2002 using GDP projections taken from the most recent OECD Economic Outlook, 1998. A similar method is used to obtain smoothed series for both commodity prices and IPD flows.

A critical element in estimating FEERs is the assumption of medium term or structural capital flows (sometimes also called the trend or target current account). Driver and Wren-Lewis, 1999a, show that FEER estimates are much more sensitive to these inputs than estimates of trend output, for example. Williamson and Mahar, 1998, provide an extensive discussion of the factors that may be involved in any off-model calculation of this input. They also provide estimates of the 'target' current account, but only for the current period, and not for the past. Farquee and Debelle,

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prices for both Japan and the US are implausibly low if the equation is estimated over the full data period.

1998, presents an econometric attempt to estimate the determinants of structural capital flows, and we have adapted their procedure.

Following Farquee and Debelle, 1998, we construct measures for equilibrium current account for our sample of G7 countries. This method allows us to investigate the role of demographics and the country's fiscal position in determining the long term saving-investment balances after controlling for business cycle fluctuations[13]. This method involves panel estimation of the following equation,

$$CAS = c_0 + c_1 CAS(-1) + c_2 (SUR - \overline{SUR}) + c_3 (DEM - \overline{DEM}) + c_4 (GAP - \overline{GAP}) \qquad (5)$$

The current account CAS is expressed as a ratio to GDP. The domestic output gap, GAP, the fiscal surplus as a share of GDP, SUR, and the demographic variable DEM[14] are expressed as deviations from their averages for the G7 countries.[15] With common coefficients across countries, and expressing the set of explanatory variables in terms of deviations from their G7 averages, this framework has the desirable property of preserving the level of the global current account in the face of global changes. The fitted equation also contains country specific fixed effects, a dummy for German Unification (from 1991:4), and a country specific dummy for the 1979 oil shock which fades out linearly over five years[16].

The common coefficients imposed across countries are significant at around the 5 per cent level of significance with the exception of the demographic variable (see Table 3).

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[12] The smoothing factor $l = 1600$ was originally chosen by Hodrick and Prescott, 1997, reflecting the cycle of US GDP and has since then become an "industry standard" for quarterly data in many applications (see, for example, Giorno *et al.*, 1995).

[13] In a larger sample of 21 countries Faruquee and Debelle, 1998, also include income per capita as a measure of the state of development. However, for the G7 countries alone, when we include this variable we find it highly statistically insignificant and wrongly signed, possibly reflecting the similarity in the stage of development across the G7.

[14] The demographic variable used by Farquee and Dabelle*,* 1998, is the age dependency ratio, the ratio of non working population over working age population. We construct a similar measure defined as the ratio of population minus labour force over labour force. For further details see the data appendix.

[15] The weights used to construct these averages are share of G7 GDP.

[16] Islam (1995) shows that as $T \to \infty$, given N, fixed effects estimation of dynamic panels is consistent and asymptotically equivalent to Maximum Likelihood Estimation.

Table 3 Panel estimates of current account equation (1977Q1 1997Q3)

Dependent Variable		*CAS*	*t-stat*
CAS(-1)		0.621	19.14
SUR		0.0017	3.39
DEM		-0.0065	-1.56
GAP		-0.0017	-4.96
adj. R^2 = 0.70	DW = 1.60	S.E.E. = 1.2	

Given the coefficient estimates, we construct the long run solution to the equation: (i.e. using CAS=CAS(-1)) as follows:

$$CAS = \frac{c_o}{1-c_1} + \frac{c_2}{1-c_1}(SUR - \overline{SUR}) + \frac{c_3}{1-c_1}(DEM - \overline{DEM}) \quad (6)$$

Note that the long run value of the output gap is set equal to zero as we are assuming internal balance[17]. The fitted values of the long run solution are reported in Figures 1-7 for the period 1973Q1 1997Q4 and are in line with those presented in Isard and Farquee, 1998.

Figures 1-7 plot three series (all as a ratio to GDP) for each of the G7: the actual current account; our estimate of structural capital flows using the method described above; and the ‘underlying’ current account. This last measure is obtained by running the model we use in the FEER calculations, but keeping the actual real exchange rate exogenous. It gives the model’s prediction for what the current account would have been if output, commodity prices and IPD had been at trend, and if the trade equations had predicted perfectly. The gap between the underlying current account and structural capital flows indicates how far the FEER has to deviate from the actual real exchange rate.

For all seven countries the actual current account series is the most erratic, and the estimated structural capital account shows least variation. The underlying current account, while smoother than the actual data, still reflects erratic movements in the actual real exchange rate. Differences between the actual and underlying current accounts usually reflect either cyclical movements in GDP, or persistent

[17] When constructing the trend current account we use filtered measures for the fiscal balances and the demographic variables to represents their longer run values.

deviations in trade volumes from their predicted levels. In the case of Canada, for example, the underlying current account in the early eighties shows a larger deficit than the actual current account. Part of this reflects a recession in Canadian GDP around this time: if output had been at trend, imports would have been higher and the current account worse. The same is true for the UK over this period. In the case of Japan at the end of the data period, our model suggests that exports were below predicted levels and imports above predicted levels, giving rise to a substantially larger trend surplus compared to the actual data.

Turning to the estimates for structural capital flows, the US trend current account is in deficit throughout most of the period, mainly reflecting demographic trends leading to a low propensity to save. The converse is true for Japan, reflecting a relatively high national saving propensity associated with a relatively low dependency ratio. The trend over time for all other countries largely reflects the smoothed movements in the structural fiscal position of the public sector for those countries. In the case of Germany, the big shift in the trend current account from significant surplus in the eighties to deficit reflects the direct impact of German unification.

In one case, Italy, the results for the first half of the data period are clearly problematic. This is a consequence of the difference between the significant pooled fiscal coefficient and the magnitude of Italy's fiscal adjustment over time. The effects of fiscal adjustment are expected to be different in a country where fiscal sustainability issues have been pronounced (Italy had a structural deficit of 10 per cent of GDP in 1992). Furthermore the unsustainability of such a large fiscal deficit implies that it would be inappropriate to regard the trend current account position as an equilibrium position. When these estimates of a sustainable current account are combined with relatively small trade elasticities (see above), we obtain quite implausible estimates for FEERs. As a result, in the remainder of the paper we concentrate on the remaining G7, and exclude Italy.

Figures 8 to 13 plot the actual and predicted effective real exchange rate for each of the G7 except Italy from 1973Q1 to 1997Q4. (In the case of Japan and France, lack of data for IPD meant that the model could not be run before 1977Q1 and 1975Q1 respectively). The real exchange rate is defined such that an increase implies a depreciation, therefore if the FEER is above the actual that implies that the real exchange rate is overvalued. We would not, of course, expect the FEER to track short term movements in the real exchange rate. However, if FEERs are to say anything about equilibrium real exchange rates, we would certainly expect trend movements in the FEER to be reflected in similar trends in the real exchange rate.

The charts suggest that in all countries the trend behaviour of the FEER is similar to the trend in the actual real exchange rate. However, there are prolonged periods when the two series diverge. For example, in the case of the UK in the early eighties and the US in the mid eighties, these currencies appear overvalued, possibly due to tight monetary policy during that period. The dollar appears to be overvalued from 1993 onwards relative to the trend values of the fundamentals, perhaps reflecting a sustained period of high economic growth. The position at the end of the data period of an overvalued dollar, and undervalued DM and Yen, is consistent with the FEER estimates derived using a different model in Driver and Wren-Lewis, 1998. We examine the correspondence between the calculated FEERs and the actual real exchange rates more formally in the next section.

4. Testing Alternative Theories of Equilibrium Exchange Rates

The existence or otherwise of a unit root in real exchange rates has important implications for both the FEER methodology and PPP. Thus, for example, if a real exchange rate does indeed contain a unit root, it means that a shock to the real rate is permanent, whereas a proponent of traditional PPP would argue that such shocks will only have a transitory effect and therefore the real rate should be mean-reverting.[18] If the real exchange rate is not mean reverting and the reasons for this are those captured by the FEER, then this should imply that deviations between the two variables will be stationary. Alternatively, if the real exchange rate is I(0), this would only be compatible with the FEER if all the variables that influence the FEER are either I(0) or cointegrated, which would seem unlikely. This section presents a series of tests for different theories of equilibrium real exchange rates, firstly by testing the time series properties of both the actual real exchange rate and the FEER separately and then by testing the relationship between them.

4.i. Testing for PPP

There are a large number of papers which test for the existence of PPP.[19] Typically the empirical evidence obtained from the application of standard univariate unit root tests suggests that PPP does not hold as they are unable to reject the null of a unit root in the real exchange rate (see, for example, Meese and Rogoff, 1988). One response to such rejection has been to say that tests for unit roots and cointegration have low power against stationary alternatives in small samples.

In an attempt to find evidence of long run PPP, researchers have turned to panel data unit root tests such as the ones developed by Levin and Lin, 1993, and Im, *et al.*, 1997. One of the main motivations for the development and application of the panel unit root tests is the increase in power compared to the well-known low power of the standard Dickey Fuller and augmented Dickey Fuller unit root tests.

The empirical evidence using these panel unit root tests tends to conflict with that cited above. Papers using panel methods report rejections of the unit root hypothesis for real exchange rates during the post Bretton Woods period (see for

[18] By traditional we mean the idea that PPP holds as a long-run equilibrium relationship, but that there may be deviations in the short run.

[19] Surveys of this literature can be found in Breuer, 1994, Froot and Rogoff, 1995, and MacDonald, 1995.

example MacDonald, 1996, Oh, 1996, Wu, 1996, Coakley and Fuertes, 1997, and Papell, 1997). However, as Papell, 1997, points out, the evidence against the unit root hypothesis is stronger for larger than for smaller panels, for monthly than for quarterly data. Additionally, by including a large number of small countries, which might be thought more likely to exhibit the property of PPP, this may explain why the G7 economies' real exchange rates appear to be non stationary when considered individually, but stationary when embodied in a larger panel. O'Connell, 1997, also finds that the real exchange rate is nonstationary when the existence of cross-sectional dependence is allowed for.

In this section, after briefly describing the testing procedures adopted, we proceed to test the PPP hypothesis using unit root techniques on the real exchange rate both at the individual country level and jointly using the panel tests developed by Im, *et al.*, 1997.

4.i.a. Panel unit root tests

The idea of increasing the span of the data using panel methods in the context of unit root testing is a relatively recent one. See Banerjee, 1999, Driver and Wren-Lewis, 1999b, and Maddala and Wu, 1999, for surveys of such tests.

Im *et al.*, 1997, propose a test procedure that tests the null hypothesis of unit roots in the series x_{it}, for all panel members (i.e $\boldsymbol{b}_i = 1$ for all i) against the alternative of stationarity for at least one panel member (i.e. $\boldsymbol{b}_i < 1$ for at least one *i*) in an equation of the form:

$$\Delta x_{it} = \boldsymbol{a}_i + \boldsymbol{d}_i t + (1 - \boldsymbol{b}_i) x_{i,t-1} + \sum_{L}^{j=1} \boldsymbol{g}_{ij} \Delta x_{i,t-j} + e_{it} \qquad (7)$$

The test proceeds by running the augmented Dickey Fuller regression (equation (7)) for each panel member and noting the t statistic, (t_{iT}) and the Lagrange Multiplier statistic (LM_{iT}) for the null hypothesis that $\boldsymbol{b}_i = 1$. After choosing an optimal lag length to ensure that the errors are serially uncorrelated, the standardised *t*-bar statistic takes the average of the individual t statistics to create an average *t*-ratio, $\bar{t}_T$ across all panel members. This is then transformed using the asymptotic values of the mean E(t_T) and variance VAR(t_T), of the average ADF statistic which Im *et al.*, 1997, tabulate for various time periods and lags, to construct a panel test which is normally distributed under the null,

$$L_{\bar{t}} = \frac{\sqrt{N}\left[\bar{t}_T - E(t_T)\right]}{\sqrt{VAR(t_T)}} \sim N(0,1) \qquad (8)$$

Im *et al.*, 1997, also provide a standardised LM-bar statistic that uses the average of the Lagrange multiplier statistics $\overline{LM}_T$ to test the null of a unit root in the panel. Again, the test statistic is transformed using the asymptotic values of the mean $E(LM_T)$ and variance $VAR(LM_T)$ such that the test statistic is also normally distributed under the null,

$$L_{\overline{LM}} = \frac{\sqrt{N}\left[\overline{LM}_T - E(LM_T)\right]}{\sqrt{VAR(LM_T)}} \sim N(0,1) \qquad (9)$$

Im *et al.*, 1997, also suggest demeaming the data to eliminate cross sectional dependence in the error terms. The demeaned series are constructed by subtracting cross section means from the observed data as follows:

$$\tilde{x}_{it} = x_{it} - \frac{1}{N}\sum_{i=1}^{N} x_{it} \qquad (10)$$

Coakley and Fuertes, 1997, suggest that this procedure should be used when the correlation matrix for the residuals from the individual augmented Dickey Fuller regressions are not diagonal.

Another popular test used in the literature is the Levin and Lin, 1993, panel tests which are based on heterogeneous panels with fixed effects. Im *et al.*, 1997, however, use stochastic simulations to show that their t-bar and LM-bar tests have substantially more power than the standard ADF tests and the Levin and Lin panel test. Maddala and Wu, 1999, also suggest that the Im *et al.*, 1997, tests have more power and exhibit better small sample properties.

4.i.b. PPP Results

As a preliminary exercise, we conduct conventional unit root tests on the individual real exchange rates series for each country for the period 1973Q1 to 1997Q4 (with the exception of France and Japan which have shorter data spans, 1975Q1 to 1997Q4 and 1977Q1 to 1997Q3 respectively). Specifically, we implement the augmented Dickey Fuller test (see Dickey and Fuller, 1979), the Weighted Symmetric test and the Phillips and Perron tests (see Phillips and Perron, 1988) and we conduct all tests both with and without a time trend. The lag length is chosen optimally using the Akaike information criteria. However, all test results are robust to choice of lag length. The results reported in Table 4 clearly show that it is not possible to reject the null of a unit root at the 5 per cent level of significance across all countries even after allowing for the inclusion of a time trend. These results are also consistent with results found by other authors (see, for example, Coakley and Fuertes, 1997).

We then apply Im *et al.*, 1997, panel unit root tests across different sample periods and all six countries. This allows us to increase the power of the test by aggregating across currencies. Maddala and Wu, 1999, suggest that the tests for Im *et al*., 1997, should be applied using a consistent lag length for all panel members. The optimal lag lengths for each country given by the AIC criteria suggest that a general lag length of $p = 4$ would be appropriate for the unit root regressions within the panel unit root test and this is imposed. However, the results reported are robust to the choice of lag length.

Initially, we present the panel unit root test results for an unbalanced panel which maximises the sample period available for all countries. The t-bar and the LM-bar results confirm that we clearly cannot reject the null of a unit root at the 5 per cent level of significance, so the real exchange series is not found to be stationary around a constant and/or a time trend (see Table 5).

Table 4 Unit Root Tests: log(RER) (1973:1-1997:4)[20]

Countries	Weighted Symmetric		Augmented Dickey Fuller		Phillips Perron	
	no trend	Trend	no trend	Trend	no trend	trend
Canada	-2.134 [0.135]	-1.981 [0.66]	-1.984 [0.293]	-2.44 [0.356]	-5.148 [0.419]	-8.067 [0.578]
Number of lags	p = 3	p = 3	p = 3	p = 3	p = 3	p = 3
France	-1.504 [0.507]	-2.388 [0.366]	-1.311 [0.624]	-2.325 [0.419]	-2.759 [0.687]	-6.314 [0.720]
Number of lags	p = 3	p = 3	p = 3	p = 3	p = 3	p = 3
Germany	-1.620 [0.421]	-2.086 [0.587]	-1.346 [0.607]	-2.126 [0.531]	-2.389 [0.733]	-7.094 [0.656]
Number of lags	p = 3	p = 3	p = 3	p =3	p = 3	p = 3
Great Britain	-0.910 [0.859]	-2.832 [0.132]	-0.941 [0.773]	-2.894 [0.163]	1.256 [0.861]	-14.18 [0.215]
Number of lags	p = 3	p = 3	p = 3	p = 3	p = 3	p = 3
Japan	-0.997 [0.824]	-2.239 [0.474]	-1.044 [0.736]	-2.711 [0.231]	-2.193 [0.757]	-9.067 [0.501]
Number of lags	p = 3	p = 3	p = 3	p = 5	p = 3	p = 5
US	-1.105 [0.772]	-1.344 [0.928]	-0.709 [0.844]	-3.285 [0.068]	-0.227 [0.944]	-14.09 [0.219]
Number of lags	p = 5	p = 5	p = 5	p = 5	p = 5	p = 5

Notes: The lag length *p* is determined using the AIC procedure. The figures in parenthesis are the *p* values.

[20] Except France (sample 1975:1 1997:4) and Japan (sample 1977:1 1997:3);

Since two of our countries (France and Japan) have a shorter data span we cannot demean the data as suggested by Im *et al*, 1997. To assess the importance of demeaning the data, rather than restrict the sample period for the majority of our countries, we drop the two countries with the short data span and estimate a four country balanced panel over the sample period 1973Q1 to 1997Q4 both using the unadjusted data and the demeaned data. Table 5 contains the results. Both the adjusted t-bar and the LM bar in both the unadjusted and demeaned data cases confirm the nonstationarity results of all the individual unit root tests and the panel unit root test conducted in the unbalanced six country panel. Furthermore, the results do not change upon the inclusion of a time trend. Hence, there is clear evidence that the real exchange rate is nonstationary and that consequently PPP does not hold for this group of countries.

Table 5 Panel Unit Root Tests: log(RER)

	Unadjusted Data				Demeaned Data			
	No trend		Trend		No Trend		Trend	
	t-bar	LM-bar	t-bar	LM-bar	t-bar	LM-bar	t-bar	LM-bar
G6 countries	0.693	-1.177	-1.242	1.235				
Smpl: unbalanced								
G4 countries*	0.433	-0.807	-1.055	1.048	-1.599	1.160	-1.104	1.082
Smpl: 1973:1 1997:4								

* France and Japan are excluded from the panel.

4.ii. Testing FEERs

Before examining the relationship between the FEER and the real exchange rate, we first analyse the statistical properties of the FEER series utilising the same testing procedures employed for the real exchange rate for each country for the period 1973Q1 to 1997Q4[21]. The results of the unit root tests for the individual FEER series are reported in Table 6 and we cannot reject the null of a unit root at the 5 per cent level of significance for any of the six counties. When we apply the panel unit root tests to all six countries, both the t-bar and the LM-bar statistics confirm that we clearly cannot reject the null of a unit root at the 5 per cent level of significance (see Table 7). Indeed, some of the test statistics are wrongly signed and therefore highly insignificant in this one tailed test. We follow the same testing procedure as before and implement the panel unit root test for the subsample of 4 countries which enjoy the largest sample period, both with and without demeaning the data. The results are consistent with the results reported above for the six countries in the unbalanced panel, except in the case where we test the demeaned data for level stationarity. In this case the LM statistic would suggest that the null can be rejected at the 5%, although not the 1%, level.

Table 7 Panel Unit Root Tests: log(FEER)

	Unadjusted Data				Demeaned Data			
	No trend		Trend		No trend		Trend	
	t-bar	LM-bar	t-bar	LM-bar	t-bar	LM-bar	t-bar	LM-bar
G6 countries	3.449	-0.517	-0.358	0.368				
Smpl: unbalanced								
G4 countries*	2.833	-0.522	-0.338	-0.342	1.378	-1.716	-0.847	1.213
Smpl: 1973:1 1997:4								

* France and Japan are excluded from the panel.

[21] Once again with the exception of France and Japan which have shorter data spans: 1975Q1 to 1997Q4, and 1977Q1 to 1997Q3 respectively.

Table 6 Unit Root Tests: log(FEER) (sample 1973:1-1997:4)[22]

Countries	Weighted Symmetric		Augmented Dickey Fuller		Phillips Perron	
	no trend	Trend	no trend	trend	no trend	trend
Canada	-1.444 [0.552]	-0.839 [0.983]	-1.858 [0.352]	-2.178 [0.502]	-3.165 [0.632]	-6.036 [0.742]
Number of lags	p = 3	p = 3	p = 3	p = 3	p = 3	p = 3
France	-0.608 [0.937]	2.187 [0.879]	-0.010 [0.957]	-3.194 [0.868]	1.853 [0.992]	-2.188 [0.965]
Number of lags	p = 4	p = 4	p = 4	p = 4	p = 4	p = 4
Germany	0.561 [0.997]	-0.876 [0.981]	1.070 [0.994]	-1.411 [0.856]	1.273 [0.993]	-3.3962 [0.920]
Number of lags	p = 2	p = 2	p = 2	p = 2	p = 2	p = 2
Great Britain	0.329 [0.995]	-1.350 [0.874]	2.263 [0.981]	-3.355 [0.920]	0.946 [0.988]	-11.97 [0.317]
Number of lags	p = 5	p = 5	p = 5	p = 5	p =5	p = 5
Japan	-0.049 [0.987]	-1.695[0.808]	0.2931 [0.977]	-2.165 [0.509]	-7.938 [0.588]	-7.938 [0.588]
Number of lags	p = 2	p = 2	p = 2	p = 2	p = 2	p = 2
US	-0.143 [0.982]	-0.056[0.984]	-1.271 [0.642]	-1.287 [0.032]	-2.149 [0.523]	-2.059 [0.773]
Number of lags	p = 4	p = 4	p = 4	p = 4	p = 4	p = 4

Notes: The lag length *p* is determined using the AIC procedure. The figures in parenthesis are the *p* values.

[22] Except France (sample 1975:1 1997:4) and Japan (sample 1977:1 1997:3);

4.iii. Testing the Relationship between FEERs and the Real Exchange Rate

Next we proceed to test whether the variable constructed by subtracting the log of the FEER from the log of the real exchange rate, henceforth DIF, is stationary,

$$DIF_{it} = \ln(FEER_{it}) - \ln(RER_{it}) \tag{11}$$

Given our unit root tests clearly demonstrate that the real exchange rate is nonstationary, the DIF variable can only be stationary if the FEER and the real exchange rate cointegrate with a cointegrating vector of (1,-1). In other words, we examine whether the I(1) behaviour of the real exchange rate can be entirely explained by similar movements in the FEER.

Once again as a preliminary exercise we conduct conventional unit root tests for each DIF series. The results presented in Table 8 provide mixed evidence of stationarity.[23] Using a 5% critical value the DIF for Germany is clearly stationary. For Canada, the UK and Japan the results depend on which test is examined as well as whether a 10% or 5% level is used. The results for the US and France appear to be nonstationary using the individual unit root tests. However, even in those countries where the DIF variable remains nonstationary in terms of *p* values it is far closer to stationarity than the real exchange rate alone. This preliminary evidence therefore suggests that the FEER is helpful in explaining trends in the real exchange rate. However, these results also suggest that, contrary to appearances in Figures 8-13, the FEER does not explain all the nonstationary behaviour of real exchange rates. In the case of France and Japan the problem appears to be located in the most recent data: if we restrict the sample to end in 1995Q4 there is evidence that the DIF variables for Japan and France are stationary (see Table 8a). For the truncated sample all the countries except Canada now show some evidence of the stationarity of the difference between the FEER and the RER.

[23] The time trend has not been included in this test. The relevant null hypothesis is that the real exchange rate and the FEER cointegrate without any divergence, as would be implied by the existence of a trend in the cointegration space.

Table 8 Unit Root Tests: $DIF_{it} = \ln(FEER_{it}) - \ln(RER_{it})$ **(1973Q1-1997Q4)**[24]

Countries	Weighted Symmetric	Augmented Dickey Fuller	Phillips Perron
Canada	-2.906* [0.015]	-2.754** [0.065]	-8.723 [0.183]
Number of lags	p = 3	p = 3	p = 3
France	-1.944 [0.216]	-1.818 [0.371]	-6.001 [0.347]
Number of lags	p = 3	p = 3	p = 3
Germany	-3.087* [0.008]	-3.082* [0.02]	-13.432* [0.058]
Number of lags	p = 2	p = 2	p = 2
Great Britain	-2.509* [0.048]	-2.394 [0.143]	-11.098** [0.092]
Number of lags	p = 3	p = 3	p = 3
Japan	-1.837 [0.276]	-1.5114 [0.527]	-10.001* [0.130]
Number of lags	p = 3	p = 3	p = 3
US	-1.819 [0.286]	-1.681 [0.440]	-10.001 [0.132]
Number of lags	p = 4	p = 4	p = 4
** Indicates that the null hypothesis can be rejected at the 10% level			
* Indicates that the null hypothesis can be rejected at the 5% level			

The lag length *p* is determined using the AIC procedure. The figures in parenthesis are the *p* values.

Table 8.A Unit Root Tests: $DIF_{it} = \ln(FEER_{it}) - \ln(RER_{it})$ **(1973Q1-1995Q4)**[25]

Countries	Weighted Symmetric	Augmented Dickey Fuller	Phillips Perron
Canada	-2.725 [0.026]	-2.559 [0.101]	-7.641 [0.231]
Number of lags	p = 3	p = 3	p = 3
France	-2.705* [0.027]	-2.786** [0.068]	-12.396** [0.075]
Number of lags	p = 3	p = 3	p = 3
Germany	-3.089* [0.008]	-3.628* [0.005]	-16.273* [0.029]
Number of lags	p = 2	p = 2	p = 2
Great Britain	-2.341**[0.077]	-2.180 [0.151]	-10.207** [0.009]
Number of lags	p = 3	p = 3	p = 3
Japan	-2.580*[0.039]	-2.394 [0.143]	-12.310** [0.077]
Number of lags	p = 3	p = 3	p = 3
US	-1.881 [0.250]	-2.044 [0.267]	-11.505** [0.093]
Number of lags	p = 4	p = 4	p = 4
** Indicates that the null hypothesis can be rejected at the 10% level			
* Indicates that the null hypothesis can be rejected at the 5% level			

The lag length *p* is determined using the AIC procedure. The figures in parenthesis are the *p* values.

These mixed results provide strong motivation for moving to the Im *et al.*, 1997, panel unit root test in order to improve the power of the individual unit root tests. The optimal lag lengths for each country given by the AIC criteria suggests that a

[24] Except France (sample 1975:1 1997:4) and Japan (sample 1977:1 1997:3);

general lag length of p = 3 would be appropriate for the unit root regressions within the panel unit root test.[26] As reported in Table 9 when we apply the panel unit root tests on all six countries, both the t-bar and the LM-bar statistics indicate that we can clearly reject the null of a unit root at the 5 per cent level of significance. We follow the same testing procedure as before and implement the panel unit root test for the 4 countries in a balanced panel and find that the adjusted t-bar and the LM-bar statistics confirm that, for both the unadjusted and demeaned data, there is clear evidence that the DIF series is stationary.[27] This is consistent with cointegration between the real exchange rate and the FEER and supports the view that the FEER methodology provides a good explanation of the determination of long run real exchange rates.

When we reduce the sample to the period 1973Q1 to 1995Q4, (see Table 9A) we find that the results for the unbalanced panel are even better while there is no material change in the results for the balanced panel, which excludes France and Japan. The reason for this is that reducing the sample period increases the stationarity of the DIF variable for France and Japan (see the discussion on individual unit root tests above).

Table 9 Panel Unit Root Tests: DIF series

	Unadjusted Data		Demeaned Data	
	t-bar	LM-bar	t-bar	LM-bar
G6 countries	-2.170	2.464		
Sample: unbalanced				
G4* countries	-2.483	3.063	-2.152	2.643
Sample: 1973:1 1997:4				

* France and Japan are excluded from the panel.

Table 9A Panel Unit Root Tests: DIF series (Data to 1995:4)

	Unadjusted Data		Demeaned Data	
	t-bar	LM-bar	t-bar	LM-bar
G6 countries	-2.979	3.602		
Sample: unbalanced				
G4* countries	-2.413	2.916	-2.001	2.293
Sample: 1973:1 1995:4				

* France and Japan are excluded from the panel.

[25] Except France (sample 1975:1 1995:4) and Japan (sample 1977:1 1995:4);

[26] Again the results reported are robust to the choice of lag length.

[27] Again a time trend has not been included for these tests. The relevant null hypothesis is that the real exchange rate and the FEER cointegrate without any divergence, as would be implied by the existence of a trend in the cointegration space.

5. Conclusion

The FEER method of calculating an equilibrium real exchange rate is the most widely used alternative to PPP, and yet FEERs have not been subjected to the same tests as PPP. In this paper we calculate historical series for FEERs for six major industrialised economies, using as far as possible the approach most commonly applied in the literature. We then ask whether movements in the FEER can explain long run movements in the real exchange rate. Specifically, we test whether real exchange rates, which for the G7 countries (excluding Italy) individually or collectively appears to be I(1), cointegrate with time series for the FEER with a coefficient of unity, so that the difference between the FEER and the real exchange rate is stationary.

Depending on the test statistic and the sample period chosen, the results at an individual country level provide some support for the FEER for all six countries. Over the full sample the results for Canada, the UK and Germany are strongest, with some support in the Japanese case. In a truncated sample the results for Japan and France are also strong. On the other hand the relationship between real exchange rate and the FEER in the US appears less significant, as the only evidence of stationarity is for the Phillips Perron test over the reduced sample. Panel unit root tests, however, suggest that the real exchange rate and FEER cointegrate. Overall therefore the results appear to suggest that, with the possible exception of US, the FEER approach does represent an improvement over PPP in explaining medium to long term trends in the real exchange rates of the major industrialised countries. It is also clear, both from these results and our failure to obtain sensible FEER estimates for Italy, that there exists scope to improve the quality of FEER estimates.

Appendix A: Data and Data Sources

Most of the series are taken from Datastream, which downloads the data from the International Financial Statistics (IFS) published by the IMF and the Quarterly National Accounts (QNA) and Monthly Economic Indicators (MEI) published by the OECD. The series are extracted for the G7 countries (Japan, US, Canada, UK, France, Italy and Germany), as well as for total OECD aggregates where appropriate. Attention is paid to ensuring that the data series for each of the countries and for the OECD, are as comparable as possible. The base period for the constant price series is 1990 and all indices are for 1990=100. The appendix contains a list of the series downloaded and the source, and the availability of the data (at the time that it was downloaded).

World variables

WT: World Trade Volume. Measure which is taken to be G7 total exports goods and services in US$ millions, deflated by implicit price deflator for total G7 exports goods and services. (1960:1 to 1997:4). Source MEI.

WT^*: Trend World Trade. Constructed using the Hodrick-Prescott Filter (with the smoothing factor $l = 1600$).

rcp: Real Commodities Prices. Constructed as a weighted average of the following commodity prices: Oil price, World food prices World beverages prices, World agricultural non food prices, World metals and minerals prices, Base metals commodity price index. Deflated by implicit price deflator for total G7 exports goods and services. (All data 1960:1 to 1997:4). Source IFS.

Individual country variables

X: Volume of Exports Goods and Services. Japan, US, UK (1960:1 to 1997:4), Canada (1961:1 to 1997:4), France (1967:1 1997:4), Italy (1970:1 to 1997:4) and Germany (1968:1 to 1997:4). National Government Sources.

M: Volume of Imports Goods and Services. Japan, US, UK (1960:1 to 1997:4), Canada (1961:1 to 1997:4), France (1967:1 1997:4), Italy (1970:1 to 1997:4) and Germany (1968:1 to 1997:4). National Government Sources

Y: Domestic Real GDP. Japan: (1960:1 to 1997:4); UK: (1960:1 to 1997:4); Italy (1970:1 to 1997:4); France: (1970:1 to 1997:4); Canada: (1961:1 to 1997:4). Germany: Pan Germany GDP from 1991, West German prior, D-Mark billion, base year 1990 (1960:1 to 1997:4). National Government Sources.

Y^*: Trend Output. Constructed using the Hodrick-Prescott Filter (with the smoothing factor $l = 1600$).

e: Real Exchange Rate. Defined as P*R/P where the foreign price level, P*, is given by the implicit price deflator for total G7 exports goods and services times the nominal dollar exchange rate (domestic/$), R, divided by domestic GDP deflator, PY. All countries (1970:1 1997:4).

rpx: Real Export Prices. Constructed by dividing the implicit price deflator for exports goods and services by domestic GDP deflator. Japan, US, UK (1960:1 to

1997:4), Canada (1961:1 to 1997:4), France (1967:1 1997:4) Italy (1970:1 to 1997:4) and Germany (1968:1 to 1997:4). Source: MEI.

rpm: Real Import Prices. Constructed by dividing the implicit price deflator for imports goods and services by domestic GDP deflator. Japan, US, UK (1960:1 to 1997:4), Canada (1961:1 to 1997:4), France (1967:1 1997:4) Italy (1970:1 to 1997:4) and Germany (1968:1 to 1997:4). Source: MEI.

A and D: External Assets and Liabilities. For all the G7 countries, except the United Kingdom, data is available from the OECD Financial Statistics Part 2, Financial Accounts of OECD Countries. For the United Kingdom, data is taken from the CSO publication, United Kingdom Balance of Payments. Annual data transformed into quarterly data by interpolation. The coverage of these publications is as follows: United States, (1960-1996), Germany, (1976-1996), Japan, (1983-1995), France, (1977-1996), Canada, (1980-1996), Italy, (1989-1996), United Kingdom, (1977-1997).

NTRAN: Net Transfers. Defined as current transfers credits minus current transfers debits, both in US dollars, converted into domestic currency using R and into real terms using PY. Data was smoothed using Hodrick-Prescott Filter. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). Source: IFS.

IIPDC: Interest Rate for IPD credit. Derived from income credit in US dollars converted into domestic currency using R and into real terms using PY. The implied interest rate on credit is calculated from quarterly estimates of A. Data was smoothed using Hodrick-Prescott Filter. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). Source: IFS.

IIPDD: Interest Rate for IPD debits. Derived from income debit in US dollars converted into domestic currency using R and into real terms using PY. The implied interest rate on debit is calculated from quarterly estimates of D. Data was smoothed using Hodrick-Prescott Filter. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). Source: IFS.

Trend Current Account Data

CAS: Current Account. Expressed as a ratio to GDP. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). Source IFS.

SUR: Fiscal Surplus. Expressed as a share of GDP. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). Source MEI.

GAP: Domestic Output Gap. All countries (1973:1 1997:4) except Japan (1977:1 1997:3) and France (1975:1 1997:4). $(Y-Y^*)$.

DEM: Demographic variable defined as the ratio of population minus labour force over labour force. Where our numerator incorporates non working age population and working age population which is not part of the labour force (i.e. not actively seeking work). Japan, US, UK, Canada, Italy (1970:1 to 1997:4) France (1975:1 1997:4) and Germany (1973:1 to 1997:4). Source: National Government Sources.

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Figure 1: US Current Account percent of GDP (1973Q1 1997Q4)

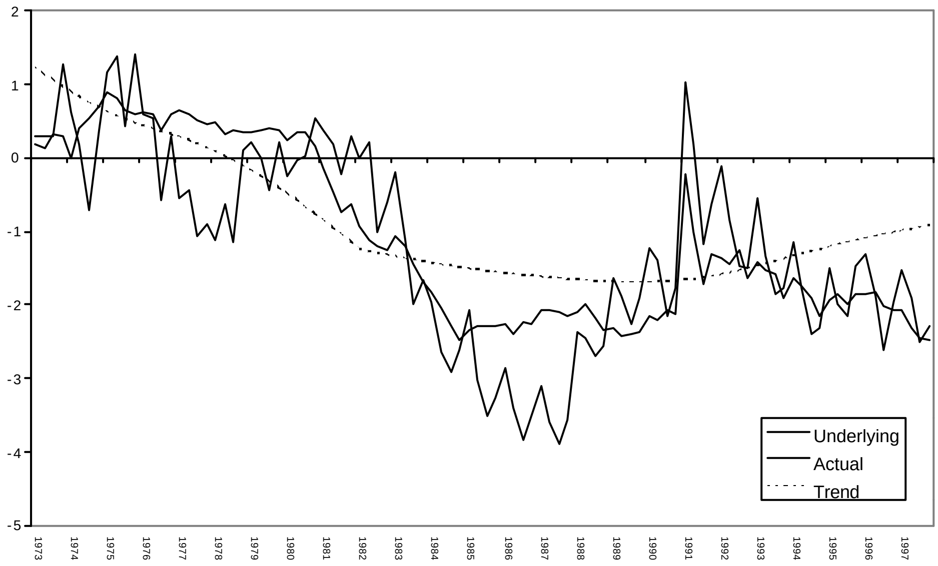

Figure 2: Japanese Current Account percent of GDP (1977Q1 1997Q3)

Figure 3: German Current Account percent of GDP (1973Q1 1997Q4)

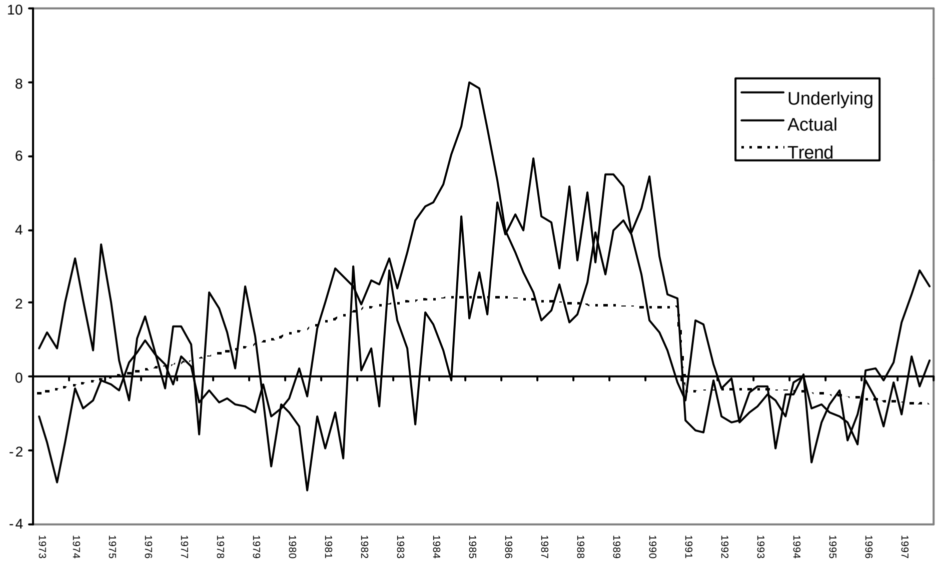

Figure 4: UK Current Account percent of GDP (1973Q1 1997Q4)

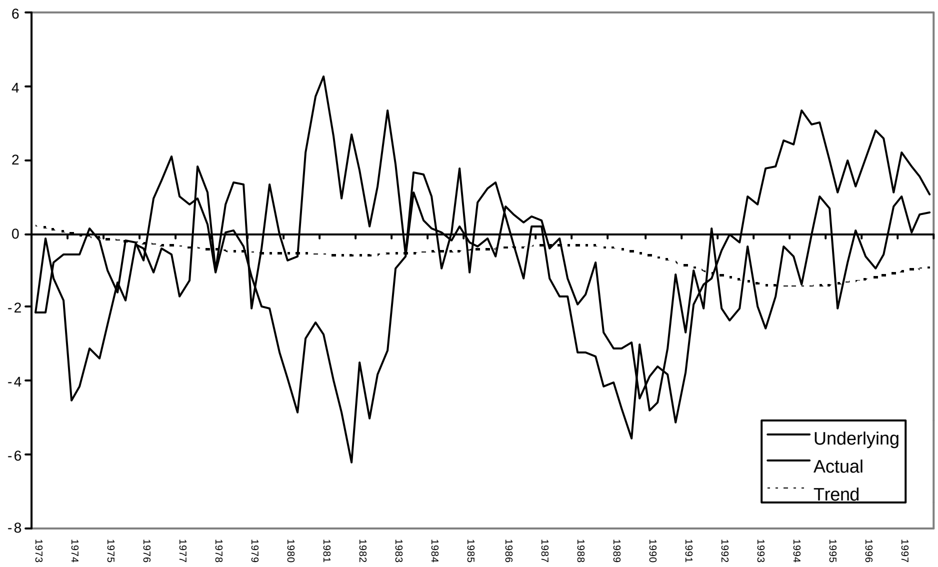

Figure 5: French Current Account percent of GDP (1975Q1 1997Q4)

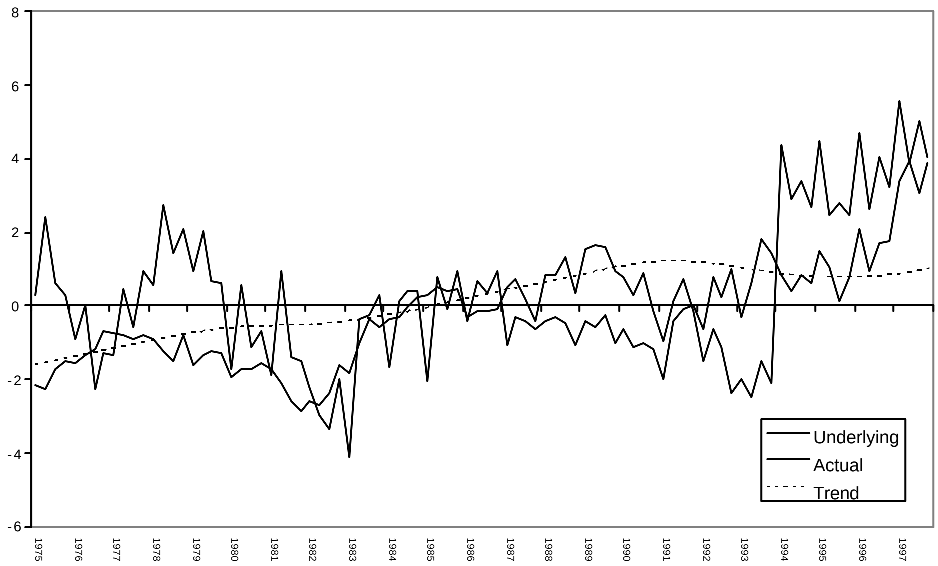

Figure 6: Canadian Current Account percent of GDP (1973Q1 1997Q4)

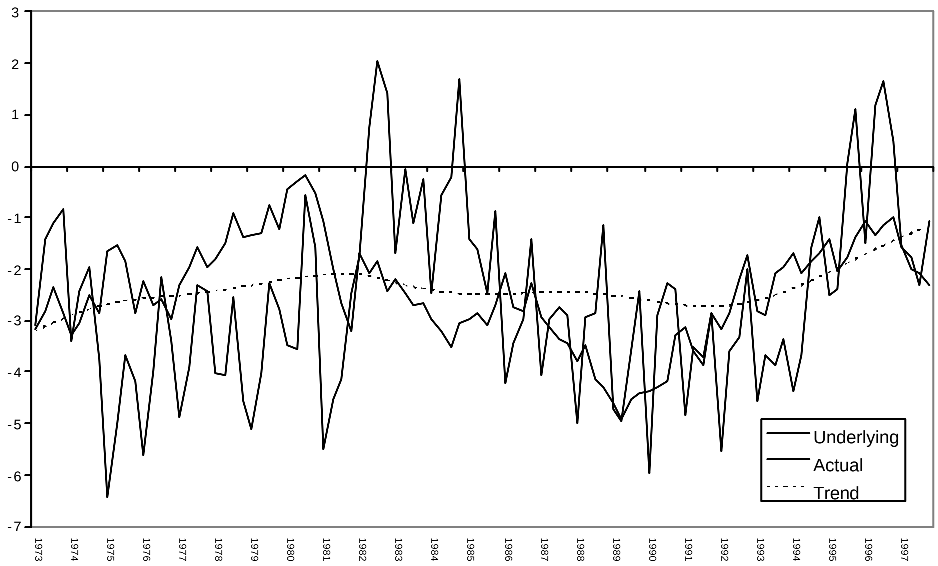

Figure 7: Italian Current Account percent of GDP (1973Q1 1997Q4)

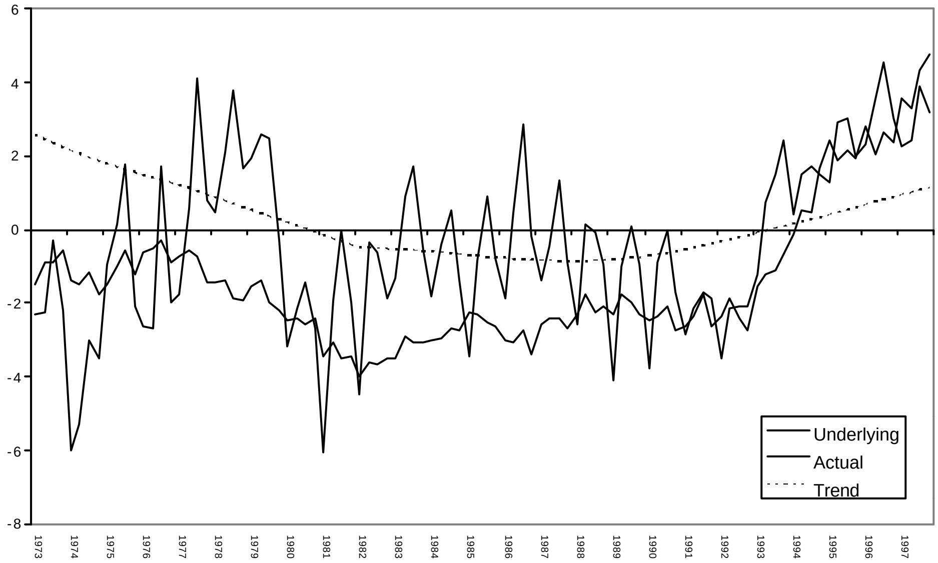

Figure 8: US Real Exchange Rate and FEER (1973Q1 1997Q4)

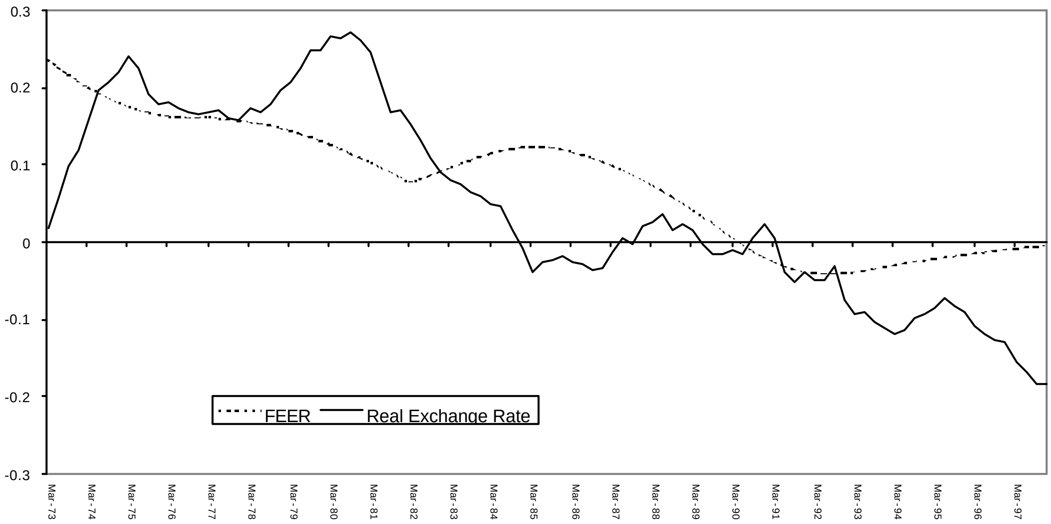

Figure 9: German Real Exchange Rate and FEER (1973Q1 1997Q4)

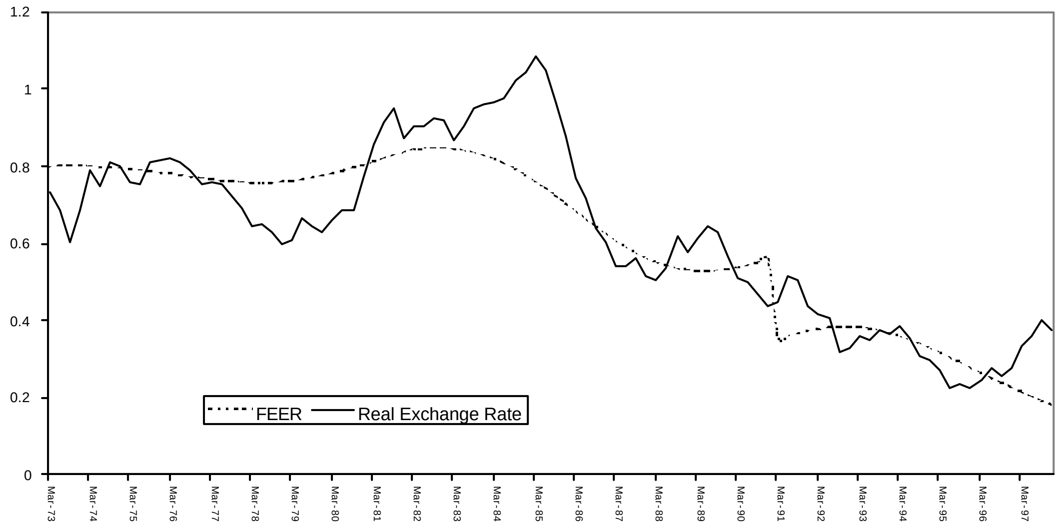

Figure 10: Japanese Real Exchange Rate and FEER (1977Q1 1997Q3)

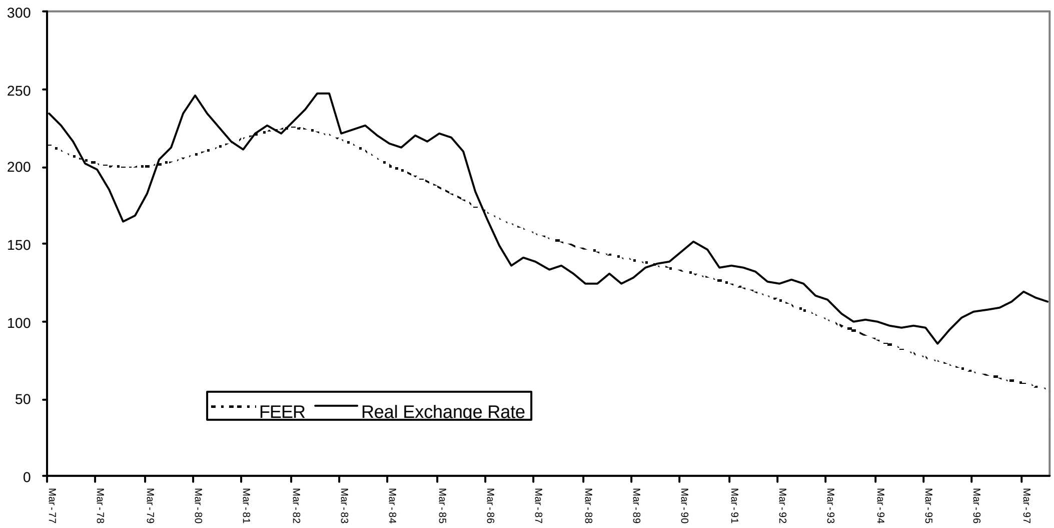

Figure 11: UK Real Exchange Rate and FEER (1973Q1 1997Q4)

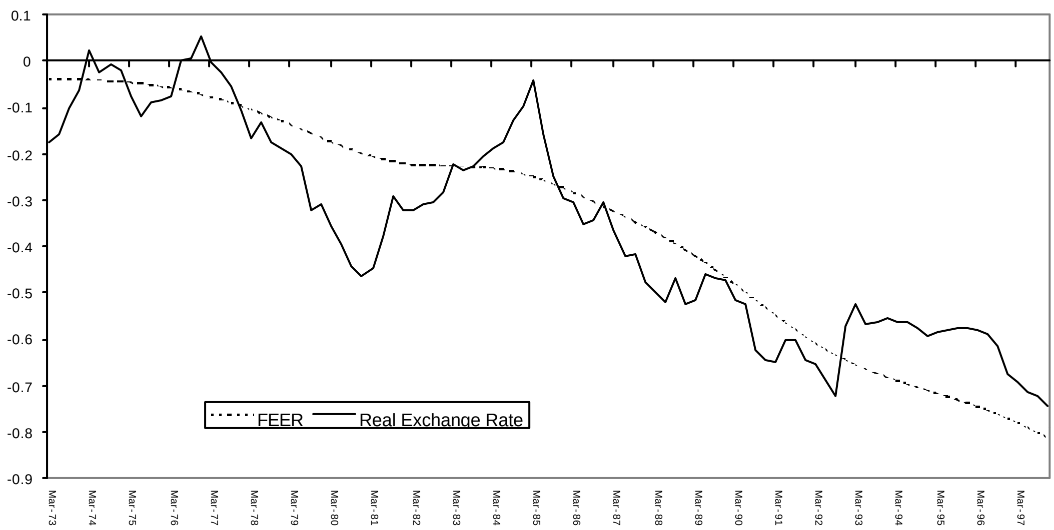

Figure 12: French Real Exchange Rate and FEER (1975Q1 1997Q4)

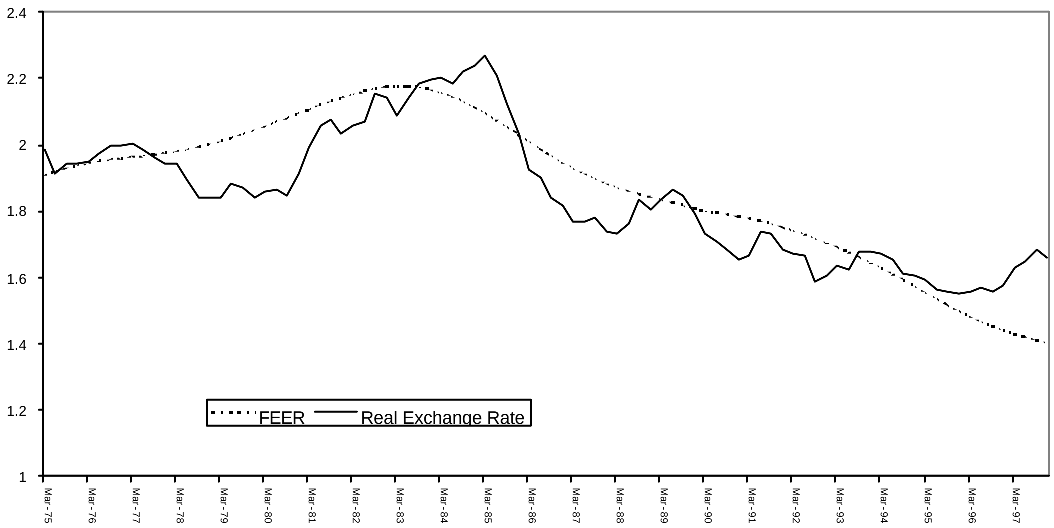

Figure 13: Canadian Real Exchange Rate and FEER (1973Q1 1997Q4)