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Discussion Papers in Economics and Econometrics

A REAL DIFFERENTIAL VIEW OF EQUILIBRIUM REAL EXCHANGE RATES

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No. 0103

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January 31, 2001

Abstract

This paper examines the interaction of G7 real exchange rates with real output and interest rate differentials. Using cointegration methods, we generally find a link between the real exchange rate and the real interest differential. This finding contrasts with the majority of the extant research on the real exchange rate - real interest rate link. We then identify a new measure of the equilibrium exchange rate in terms of the permanent component of the real exchange rate that is consistent with the dynamic equilibrium given by the cointegrating relation. Furthermore, the presence of cointegration also allows us to identify real, nominal and transitory disturbances with only minimal identifying restrictions. Our findings suggest that misalignments are largely due to nominal shocks, but that their half-life is much lower than is suggested when purchasing power parity is used as the reference equilibrium. This has important implications for the persistence measures of real exchange rates that are reported elsewhere in the literature.

Keywords: Equilibrium Exchange Rates; Cointegration; Permanent and Transitory Decomposition.

JEL Classification: F31

*A substantial part of this paper was prepared while Hoffmann was visiting the Center for Financial Studies (CFS) at Johann-Wolfgang von Goethe University (Frankfurt/Main). The hospitality of the CFS is gratefully acknowledged. We would also like to thank seminar participants at Frankfurt and at the Danish Social Science Research Council Conference on the Macroeconomic Transmission Mechanism in Copenhagen, in particular Hans-Christian Kongsted, Axel Weber and Nuno Casolla.

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1 Introduction

Recent policy-based discussions on the causes of the strength of the US dollar against the euro have focussed on the importance of two key real differentials between the euro-zone and the US, namely a real yield differential and the relative real GDP growth differential (Corsetti (2000)). Indeed, many well-known exchange rate models would highlight these two differentials as being key determinants of real exchange rates. For example, sticky price models (see Dornbusch (1976) and Mussa (1984)) and optimising models (see, for example, Grilli and Roubini (1992)) emphasize the effect of liquidity impulses on real interest rates and the real exchange rate. The Balassa-Samuelson effect would predict that relatively fast real GDP growth in a country should produce an appreciation of its real exchange rate, to the extent that relatively rapid productivity growth is positively correlated with relative GDP growth. In this paper we attempt to make a contribution to the empirical exchange rate literature by analysing the real effective exchange rates of the G7 countries in terms of these two key differentials. Our approach involves taking a simple VAR representation of the real exchange rate and real interest rate and GDP differentials and uses this to unravel the relative explanatory power of real, nominal and transitory shocks for the G7 real effective exchange rates, over the period 1978 quarter 2 to 1997, quarter 4. Our contributions relate to the estimation and specification of the variables entering the VAR, the way real and nominal shocks are identified from such a VAR and the appropriate way to measure exchange rate misalignment.

Our approach is in the spirit of the literature started by Clarida and Gali (1994), which we label the SVAR approach to real exchange rate modelling (see also Weber (1997) and Rogers (1999)). In essence, the SVAR literature focuses on the relative importance of demand, supply and nominal shocks in explaining real exchange rate movements. However, practically all of the papers in this literature ignore a variable which is pivotal to most exchange rate models, namely the real interest differential. Furthermore, the common practice of using a vector autoregressive model in first differences means that any potential cointegrating relationships amongst the variables driving the shocks are ignored and it is conceivable that the introduction of the levels of the variables has implications for the findings reported by others. Using cointegrated vector autoregressions, we find a long-term link between real differentials and real exchange rates and we provide a theoretical rationale for this.

The perception that exchange rates can spend long periods away from fundamentals-based measures of equilibrium has led to a revival of interest in the concept of exchange rate misalignment. In the SVAR literature, measures of the equilibrium exchange rate are often constructed from the estimated VAR, and in our work we define the equilibrium real exchange rate as that which is consistent with the dynamic equilibrium represented by the cointegrating relationship found in the data. We compare this cointegration-based notion of an equilibrium exchange rate to that proposed by Clarida and Gali (1994). It turns out that the cointegration based method delivers measures of misalignment that are highly correlated with a Clarida-Gali type measure of misalignment, based only on real shocks.[1] However, while the level of the real exchange rate seems to be largely determined by real factors, we find that nominal forces play an important role in the variability of its changes at all forecast horizons. Our results may also have important implications for the discussion about the persistence of real exchange rates, the so-called PPP puzzle (Rogoff (1995)). It turns out that the misalignments we identify are much less persistent than the ones obtained from studies where the notion of the equilibrium real exchange rate is based on purchasing power parity.

The findings of this paper also seem to build a bridge between some conflicting results in the literature. For example, earlier contributions (e.g. Clarida and Gali (1994), Weber (1997) and Rogers (1999)) tended to interpret real shocks as permanent and nominal shocks as transitory and work with a first difference representation of a VAR. Depending on the information set used, these contributions would either emphasize the role of nominal disturbances (such as Clarida and Gali (1994) and Rogers (1999)) or find a very important role for permanent disturbances (such as Weber (1997)). Our approach identifies permanent and transitory components based only on the cointegrating information in the data. We then disentangle real and nominal permanent shocks using the approach pioneered by Blanchard and Quah (1989). We find that the level of the real exchange rate is mainly explained by permanent shocks but that, at least in G7 data, both real and nominal forces play an important role for higher-frequency variability.

The remainder of the paper is structured as follows. In Section two we provide a brief review of the literature that aims to explain exchange rate persistence and volatility. In Section three we provide a motivational framework for our empirical analysis, while in section four the identification of permanent and transitory shocks is described. The data is described in Section 5 and our econometric model and empirical results are presented in Section 6. Section Seven concludes.

[1] See MacDonald (2000) for an overview of the literature on exchange rate misalignment.

2 Exchange Rate Persistence and Volatility: A Motivational Overview.

In this section we present a brief overview of some of the key empirical results concerning the persistence and variability of real exchange rates which have a bearing on the empirical results presented in section 4. A useful starting point is to define the real exchange rate, q_t, as:

$$q_t = s_t - p_t + p_t^*, \tag{1}$$

where s_t denotes the nominal exchange rate (home currency price of a unit of foreign exchange), p_t denotes a price level, an asterisk denotes a foreign magnitude and lower case letters denote that a logarithmic transformation has been used. Although a strict interpretation of PPP requires that the logarithm of the real exchange rate be zero (or equal to a constant term in the presence of non-zero transaction costs or due to the use of price indices) most proponents of PPP would regard a rapidly mean-reverting real exchange rate as supportive of PPP. In particular, a half-life mean reversion speed of around one year is normally taken to be the benchmark number consistent with PPP (see Rogoff (1995)).

Using univariate unit root methods and data for the recent floating period a number of researchers find that real exchange rates are effectively unit root process and do not exhibit any significant mean reversion (see Rogoff (1996) and MacDonald (1995a)). However, a significant transitory, or mean reversion, component is recovered from long time span data sets (see, for example, Edison (1987), Frankel (1986,1988), Abuaf and Jorion (1990), Grilli and Kaminski (1991) and Diebold, Husted and Rush (1991), Cheung and Lai (1994))[2] and in studies which exploit panel data sets for the recent floating period, (see, *inter alia*, Frankel and Rose (1995), Wu (1995), Oh (1995) and MacDonald (1995b)). The typical half-life in these studies is between 3 and 5 years and this is regarded as being too slow to be consistent with PPP, hence the label the PPP puzzle.[3]

[2] Engel (2000) has demonstrated that there can be large size biases in tests for long-run PPP and, indeed, that there may be a significant unit root component that is not detected by these tests. He associates this non-stationary component with secular movements in the relative price of non-traded goods.

[3] On the basis of an impulse response analysis, Cheung and Lai (2000a) argue that the confidence intervals of half-life estimates are generally rather wide, and this suggests that a researcher should be cautious in accepting extant point estimates of half-lives as a precise measure of parity reversion. Furthermore, Cheung and Lai (2000b) emphasize that since both panel and long horizon data sets focus on industrial countries there may be issues of survivorship bias (Froot and Rogoff (1995); that is, real exchange rates are more likely to

How may the PPP puzzle be explained? One explanation, which is in the spirit of traditional PPP, involves recognizing the implications that non-zero transaction costs can have for the time series properties of real exchange rates. In particular, a number of theoretical papers (see, for example, Dumas (1992) and Sercu, Uppal and Van Hulle (1995)) have demonstrated that if markets are spatially separate, and feature proportional transactions costs, deviations from PPP should follow a non-linear mean-reverting process, with the speed of mean reversion depending on the magnitude of the deviation from PPP. A number of papers (see, inter alia, Obstfeld and Taylor (1997), Michael, Nobay and Peel (1997), O'Connell (1996) and O'Connell and Wei (1997)) have implemented this idea using band threshold autoregressive models and find support for the non-linear hypothesis and report half-lives which are much closer in spirit to a traditional form of PPP.[4]

An alternative way of reconciling these findings is to consider the pricing-to-market and pass-through models that have received increased attention in the recent so-called 'new open-economy' macroeconomics (see Lane (1999) for a survey). Under the pricing to market (PTM) scheme, producers engage in price discrimination by setting different markups over marginal costs in domestic and foreign markets and adjust the markup to absorb nominal exchange rate changes. The existence of PTM has been used to explain both deviations from PPP (as in Feenstra and Kendall (1997)) and the excessive volatility of real exchange rates (as in Betts and Devereux (1996)). In recent work, however, Obstfeld and Rogoff (1999) cast doubt on the empirical relevance of the PTM-LCP assumption, in spite of its ability to help in rationalizing key facts of international macroeconomic fluctuations. One reason for this is to be found in the work of Rogers and Jenkins (1995) and Wei and Parsley (1995). These researchers show that adjustment speeds for disaggregate prices are similar to the adjustment speeds found for aggregate CPI real exchange rates and this seems inconsistent, at least, with the PTM story since it would imply that their is a one-to-one relationship between the firms pricing policy and the exchange rate.

An alternative explanation for real exchange rate persistence involves rec-

be stationary for such countries because Balassa-Samuelson productivity type effects are unlikely to be very strong compared to the real exchange rates of developing countries. However, using real exchange rate data for 94 countries, Cheung and Lai (2000b) report the opposite result - parity reversion is less strong in industrial countries than in developing countries.

[4]For example, Obstfeld and Taylor use a band threshold autoregressive model to estimate mean reversion speeds for real exchange rates, defined using both CPI and disaggregate price series. For the CPI-based real exchange rates they report adjustment speeds outside the transaction band of one year, while for the disaggregate prices they report adjustment speeds as low as 2 months.

ognizing that there are real determinants of real exchange rates, such as net foreign assets and Balassa-Samuelson productivity effects. These may be motivated using the so-called real exchange rate - real interest rate differential:

$$q_t = \bar{q}_t - \varphi(r_t - r_t^*), \tag{2}$$

where $\bar{q}_t$denotes the long-run real exchange rate and $r_t - r_t^*$ is the real interest differential. This relationship may be derived by exploiting UIP, the fisher closed conditions and ex ante PPP (as in Meese and Rogoff (1984) and Edison and Melick (1995)) or, as in Obstfeld and Rogoff (1996), using UIP, the Fisher closed conditions and a Phillips curve relationship. In the former derivation, φ is a function of the maturity of the bonds underpinning the real interest rates, while in the latter interpretation it is a function of parameters from the Phillips curve relationship. In estimating (2) a number of papers have simply assumed $\bar{q}_t$to be constant (i.e. $\bar{q}_t = \alpha$). This strand of research generally finds an absence of a cointegrating relationship for the vector implied by (5) when the Engle-Granger two-step method is used (see, inter alia, Meese and Rogoff (1984) and Edison and Melick (1995), Throop (1994) and Coughlin and Koedijk (1990)), but somewhat stronger evidence when the maximum likelihood estimator of Johansen is employed (see, inter alia, Edison and Melick (1992,1995), MacDonald (1997)). Studies which model $\bar{q}_t$as a function of 'real' fundamentals such as net foreign assets and Balassa-Samuelson effects generally find more favourable long-run cointegrating relationships (see Chinn and Johnston (1999), Gagnon (1996) and Lane and Miles-Ferreti (2000)).

The permanence of real exchange rates has been addressed in a somewhat separate strand of the empirical exchange rate literature. In particular, a number of researchers have used both univariate and multivariate Beveridge-Nelson decompositions to decompose real exchange rates into permanent and transitory components (see, for example, Huizinga (1987), Cumby and Huizinga (1990), Clarida and Gali (1994) and Baxter (1994)):

$$q_t = q_t^P + q_t^T, \tag{3}$$

where q_t^P and q_t^T are the permanent and transitory components of the real exchange rate, respectively. The general tenor of these results is that when a univariate decomposition is used the permanent component of the real exchange rate is around 0.9 (see, for example, Huizinga (1987)), but when a multivariate decomposition is used the split between the permanent and transitory components is more evenly balanced (see, for example, Clarida and Gali (1994)). These results would therefore seem to reinforce the unit root results: real exchange rates although persistent seem to contain important mean-reverting, or transitory, elements. Both Huizinga (1987) and Clarida and Gali (1994) use the permanent component of the real exchange rate as a measure of the equilibrium real exchange rate and the gap between the actual and permanent is the extent of misalignment.

A somewhat different approach to decomposing the permanent and temporary components of the real exchange rate has been advocated by Clarida and Gali (1994). In particular, they consider the vector:

$$\Delta x_t' = [\Delta y_t, \Delta q_t, \pi_t], \tag{4}$$

where y_t, denotes relative output (home-foreign) and π_t denotes relative inflation. Using a trivariate VAR modeling approach and the identification methods of Blanchard and Quah on the long-run, C(1), matrix, Clarida and Gali are able to identify three shocks from this vector: a supply shock, a demand shock and a nominal shock.[5] Using this framework Clarida and Gali produce variance decompositions of the real US dollar bilateral rates of the Canadian dollar, German mark, Japanese yen and UK pound for the recent floating period. Of this total, almost all is attributable to demand shocks in the case of the UK and Canada, while for Japan the split is 60% demand and 30 % monetary, whith the split being approximately equal for the German mark. The proportion of the forecast error variance due to the supply shock is statistically insignificant at all forecast horizons. The very small supply side specific component reported by Clarida and Gali has been confirmed by others (see, for example, Chadha and Prasad (1997)) for different currencies and different time periods and has indeed become something of stylized fact in the literature on the economics of real exchange rates. However, one reason why Clarida and Gali find such a small supply side component may reflect the actual specification of the supply side used in their model. Both Rogers (1995) and Weber (1998) have reworked the Clarida and Gali analysis using a richer supply side specification and find that the supply side, or permanent component of the real exchange rate puts in a much more respectable showing of approximately 30 per cent. Clarida and Gali (1994) have proposed using the real exchange rate with the nominal shock netted

[5] The particular identifying restrictions used (based on a modified version of the Mundell-Fleming-Dornbusch (MFD) model) are: money, or nominal, shocks do not influence the real exchange rate or relative output in the long run; only supply shocks are expected to influence relative output levels in the long run; both supply and demand shocks are expected to influence the real exchange rate in the long-run.

out as their measure of the permanent, or equilibrium, exchange rate - an exchange rate misalignment is therefore determined by the nominal shock.

We summarize this section by noting that, on the basis of a number of different approaches, real exchange rates are highly persistent. Various interpretations have been proposed to explain this persistence, such as recognizing non-linearities in real exchange rates, the importance of pricing to market policies of companies and the relationship between the persistence in real fundamentals and real exchange rates. In this paper we pursue the latter interpretation in the context of a structured VAR approach. In contrast to previous estimates of such VARs we make, at least, two contributions. First, we explicitly condition the real exchange rate on a real interest differential, a variable which has been ignored in previous structured VARs and, secondly, we explicitly recognize the potential long-run or cointegrating relationships amongst the variables entering the VAR. Our estimated SVAR is then used to shed further light on the sources of real exchange rate variability discussed in this section.

3 Real Exchange Rates and Real Differentials

As we have indicated, the explanation adopted in this paper for real exchange rate behaviour relies on a real - nominal decomposition. A useful starting point for such a decomposition is equation (5), which we repeat here:

$$q_t = \bar{q}_t - \varphi(r_t - r_t^*). \tag{5}$$

What are the factors driving $\bar{q}_t$likely to be? Given that we want our VAR model to be highly parsimonious, a useful way of thinking about this question is to consider the relationship between the real exchange rate and per capita GDP which has been documented in a number of papers. For example, Kravis and Lispsey (1983, 1987, 1988) have demonstrated that a negative correlation between the real exchange rate (as defined here) and per capita real gross domestic product is robust across numerous cross-sectional specifications. Furthermore, Bergstrand (1991), using Kravis and Lipsey data, illustrates that 87 per cent of the variation in real exchange rates of 21 countries in 1975 is expained by per capita GDP and a constant. This of course is very much a black box relationship, in the sense that it is not informative about what causes the relationship. A useful way of thinking about the sources of the relationship between a GDP differential and the real exchange rate is to consider the familiar decomposition of the overall real exchange rate into the relative price of traded goods and the internal price ratio. If we assume that the general prices entering our definition of the real exchange rate, (1), can be decomposed into traded and non-traded components as:

$$p_t = \alpha_t p_t^{NT} + (1 - \alpha_t) p_t^T, \tag{6}$$

$$p_t^* = \alpha_t p_t^{NT*} + (1 - \alpha_t) p_t^{T*}, \tag{7}$$

where p_t^T denotes the price of traded goods, p_t^{NT} denotes the price of non-traded goods and the α_t's denote the share of non-traded goods in the overall price level (and are assumed to be the same across countries). Additionally, assume that a similar relationship to (1) can be defined for traded goods as:

$$q_t^T = s_t - p_t^T + p_t^{T*}. \tag{8}$$

By substituting (6), (7) and (8) in (1) the following expression may be obtained:

$$q_t = q_t^T + [\alpha(p_t^{NT*} - p_t^{T*}) - \alpha(p_t^{NT} - p_t^T)], \tag{9}$$

$$q_t = q_t^T + q_t^{T,NT}, \tag{10}$$

$$q_t^{NT,T} = \alpha[(p_t^{NT*} - p_t^{T*}) - (p_t^{NT} - p_t^T)]. \tag{11}$$

The first term in (9), q_t^T, represents the law of one price (LOOP), or violations of the LOOP, while the second term, $q_t^{NT,T}$, represents the so-called internal relative price ratio.

According to the Balassa-Samuelson hypothesis, the LOOP is assumed to hold continuously and relatively rich countries - defined as countries with a relatively high per capita GDP - have absolute productivity advantages in the production of both traded and non-traded goods, but a relative productivity advantage in traded goods compared to their trading partner(s). With some mechanism equalising wages within countries, the relative price of non-traded goods will be higher in the country with the larger per capita income. If the home country is the country with the high per capita income, expression (9) predicts it will have an appreciated real exchange rate, defined using overall prices.

A second supply side influence on the internal price ratio involves relative factor endowments. In the traditional Hecksher-Ohlin two factor, two good, relative factor endowments model, nontraded (traded) goods are asssumed to be relatively labour-intensive (capital-intensive) in production. High per capita income countries are assumed to have a comparative advantage in producing traded goods and so the the relative price of non-traded goods will be higher in countries with relatively high per capita income.

In addition to these supply side influences, there is also likely to be a demand side effect on the internal price ratio. For example, both Dornbusch (1988) and Neary (1988) note that changes in tastes could produce changes in the relative internal price ratio which have a similar affect on the real exchange rate to the Balassa-Samuelson effect. Bergstrand (1991) formalises a proposition by Linder (1961) that per capita income is likely to be the most important single determinant of the demand structure within a country. In particular, Bergstrand uses a nonhomethetic nested Cobb-Douglas-Stone-Geary utility function for the representative consumer worker of the following form:

$$u_t = (x^T - \overline{x}^T)^\delta (x^{NT} - \overline{x}^{NT})^{1-\delta},$$

where x^T (x^{NT}) denotes the amount consumed of the traded (nontraded) commodity and an overbar denotes the exogenous mimimum-consumption requirement. Maximising this relationship subject to a standard budget constraint produces a demand relationship for the non-traded good relative to the traded good in which a one percent rise in per capita GDP in the home country will cause the home country's relative demand for the nontraded good to be higher than the foreign country, if the weighted minimum-consumption requirement for the traded commodity is greater than that for the nontraded sector.[6]

Bergstrand seeks to unravel the seperate roles of these three different explanations for the real exchange rate - per capita real GDP relationship using a cross section of 23 real exchange rates in 1975. He shows that all three effects are statisically significant determinants of real exchange rates. In this paper we do not seek to seperate the influence of these seperate sources on the

[6]The idea being that for a country with a relatively low level of per capita GDP the minimum consumption requirement for commodities is likely to be greater than that for services.

real exchange rates. Rather we assume that they are subsumed within our measure of per capita real income and focus on this as the key determinant of $\bar{q}_t$.

In the light of the above discussion, we rewrite equation (5) in the following, more general, way:

$$q_t = \beta_1 (y_t - y_t^*) + \beta_2 (r_t - r_t^*), \qquad \beta_1, \beta_2 < 0. \tag{12}$$

The second novelty we introduce in our modelling is that we do not require q and $\overline{q}$ to cointegrate. Rather, we allow for the non-stationarity of the real interest rate differential. Even though some economists would certainly question the non-stationarity of real interest rates, we would like to emphasize that the notion of non-stationarity, and hence of a long-run that we employ throughout this paper, is sample dependent. However, the data do not allow us to impose stationarity on the real exchange rate, even when full-information system methods are used. Furthermore, recent theoretical and empirical research documents that - at least over the sample period that we are going to investigate - real interest rates tend to follow very persistent processes. For example, many studies show that the *de-facto* behaviour of central banks can be well approximated by a real interest rate rule (see e.g. Romer (2000) and Rotondi (2000) and the literature surveyed there for theoretical expositions and Clarida, Gali and Gertler (1998) for empirical evidence). If real interest rates are generated by such a rule, they may become very persistent and virtually indistinguishable from an $I(1)$ process.[7]

Against this backdrop, equation (12) suggests a cointegrating relationship between real differentials and the real exchange rate. As we will show, G7 data allow us to establish this relationship. This suggests that another real differential, i.e. the GDP differential, may have been missing in the search for the link between real exchange rates and real interest rate differentials.

In order to identify the cointegrating relation suggested by (12), we used a cointegrated VAR, or vector error-correction model, of the form:

$$\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{X}_{t-1} + \boldsymbol{\varepsilon}_t, \tag{13}$$

where $\mathbf{X}_t = \left[\begin{array}{ccc} (y - y^*), & q, & (r - r^*) \end{array}\right]_t'$. The parameters of the model are $\Gamma(\mathbf{L})$ which is a 3×3 matrix polynomial in the lag operator $\mathbf{L}$ and the matrix of cointegrating vectors $\boldsymbol{\beta}$ and the error-correction loading matrix $\boldsymbol{\alpha}$ as well as the *i.i.d.* disturbance vector $\boldsymbol{\varepsilon}$, with covariance-matrix Ω. The VECM can be inverted to produce a moving average representation of the following form:

[7] King, Plosser, Stock and Watson (1991) find some evidence of I(1) ness for US real interest rates.

$$\mathbf{X}_t = \mathbf{A}\sum_{i=1}^{t}\boldsymbol{\varepsilon}_i + \mathbf{A}^*(\mathbf{L})\boldsymbol{\varepsilon}_t, \tag{14}$$

where the first terms is the random walk, or permanent, component of $\mathbf{X}$ and the second term is a stationary moving average. Johansen (1995) demonstrates that:

$$\mathbf{A} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}_{\perp}^{'}\Gamma(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}_{\perp}^{'}, \tag{15}$$

and $\boldsymbol{\beta}_{\perp}$ and $\boldsymbol{\alpha}_{\perp}^{'}$ are the orthogonal complements to $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$, respectively. It is useful to rewrite (14) as:

$$\mathbf{X}_t = \mathbf{A}_0\sum_{i=1}^{t}\boldsymbol{\pi}_i + \mathbf{A}^*(\mathbf{L})\boldsymbol{\varepsilon}_t, \tag{16}$$

where $\boldsymbol{\pi}$, the vector of permanent shocks, is a linear combination of the reduced form residuals given by:

$$\boldsymbol{\pi} = \alpha_{\perp}^{'}\varepsilon_t \tag{17}$$

and from Johansen (1995) the loadings matrix, $\mathbf{A}_0$ is given by:

$$\mathbf{A}_0 = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}_{\perp}^{'}\Gamma(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}. \tag{18}$$

In our three-dimensional system with one cointegrating relationship, $\mathbf{A}_0$ is 3×2 and $\boldsymbol{\pi}_t$ is 2×1.

4 Shock identification and equilibrium exchange rates

In this section, we will contrast two approaches to the measurement of equilibrium exchange rates. The first is based on cointegration information in the data. Even though the cointegrating relationship we identify has a theoretical interpretation, we will not need further information to extract a permanent component from our multivariate information set. In this respect, our approach is relatively atheoretical and we will refer to it as the P-T approach, because it exploits recent developments in permanent-transitory decompositions in cointegrated systems.

The second approach we use is rooted in the structural VAR literature and as such relies on identifying assumptions that are provided by economic theory. We will refer to it as the 'Clarida-Gali' approach. This method identifies structurally meaningful shocks, some of which may be deemed fundamental, whereas others are supposed to be due to non-fundamental disturbances. Once these shocks have been identified from the system, it is possible to reconstruct the time-series components, say of the real exchange rate, that are explained by both fundamental and non-fundamental shocks. This provides us with a measure both of the equilibrium exchange rate and misalignment over the sample period.

4.1 The PT approach

The recent literature on cointegrated systems shows that the transitory part of a multivariate time-series can be expressed as a linear combination of the deviation of the cointegrating relationships from their mean, i.e. the cointegration or equilibrium error (see Gonzalo and Granger (1995), Proietti (1997) and Johansen (1997)). In this paper, we use Johansen's (1997) modification of the Gonzalo-Granger decomposition:

$$\mathbf{X}_t = \mathbf{A}(\mathbf{1})\Gamma(\mathbf{1})\mathbf{X}_t + \left[\mathbf{I} - \mathbf{A}(\mathbf{1})\Gamma(\mathbf{1})\right]\mathbf{X}_t. \tag{19}$$

Proietti (1997) has shown that

$$\left[\mathbf{I} - \mathbf{A}(\mathbf{1})\Gamma(\mathbf{1})\right]\mathbf{X}_t = \boldsymbol{\psi}\boldsymbol{\beta}'\mathbf{X}_t$$

which implies that - very much as in the Gonzalo-Granger decomposition - the transitory part of $\mathbf{X}_t$ is a linear combination of the cointegrating error.

In the context of this paper, the cointegration error, $\boldsymbol{\beta}'\mathbf{X}_t$, measures the deviation of the data from the steady-state relationship (12) and as such represents a natural measure of exchange rate misalignment.

4.2 The CG-approach

An alternative method to calculate an equilibrium exchange rate is based on the decomposition of the involved time series into historical shock components. Clarida and Gali (1994) were the first to apply this technique to real exchange rate data and we therefore refer to this approach as the CG method. Clarida and Gali decompose the real exchange rate into two components: the permanent part that is attributed to real shocks and a transitory or 'misalignment' component that is related to nominal disturbances.

Even though our model is similar to Clarida and Gali's (1994) in that we are dealing with a three-variable setup, the informational requirements for just identification are reduced due to the presence of cointegration; in fact, the presence of cointegration enables us to just-identify our model using a version of Blanchard-Quah's (1989) procedure. To see this, note that from our discussion in Section 3 the permanent component of X_t can be written as:

$$\mathbf{X}_t^P = \mathbf{A}_0 \sum_{l=0}^{t} \boldsymbol{\pi}_l,$$

where, as we have seen, $\mathbf{A}_0$ is a function of the $\boldsymbol{\beta}_\perp$ and $\boldsymbol{\alpha}_\perp$ vectors. The latter, however, only determined up to a linear transformation. Hence, $\mathbf{X}_t^P$ remains the same whenever we choose $\widetilde{\pi}_t = \mathbf{S}\boldsymbol{\pi}_t$ and $\widetilde{\mathbf{A}_0} = \mathbf{A}_0\mathbf{S}^{-1}$ for any non-singular 2×2-matrix $\mathbf{S}$. So, for any initial choice of $\boldsymbol{\beta}_\perp$ and $\boldsymbol{\alpha}_\perp$, how should we choose $\mathbf{S}$?

We start by requiring that the two permanent shocks be orthogonal and have unit variance. Hence, we get:

$$var(\pi_t) = \mathbf{S}\boldsymbol{\alpha}_\perp' \Omega \boldsymbol{\alpha}_\perp \mathbf{S}' = \mathbf{I}_2. \tag{20}$$

This gives us three non-redundant restrictions on $\mathbf{S}$. The fourth restriction that is required to just-identify the four elements of $\mathbf{S}$ comes from the theoretical model: in the presence of $I(1)$ real interest rates, money shocks can have a 'long-run' impact on the real exchange rate but not on the output differential. Requiring the first of the two permanent shocks to be the supply shock and the second the money (nominal) shock and bearing in mind the ordering of variables in $\mathbf{X}_t$, this amounts to requiring that:

$$\widetilde{\mathbf{A}_0} = \begin{bmatrix} a_{11} & 0 \\ a_{21} & a_{22} \\ a_{31} & a_{32} \end{bmatrix} = \mathbf{A}_0\mathbf{S}^{-1}. \tag{21}$$

This completes the identification of the permanent shocks. To just identify our model, we also need to identify a third shock which will be purely transitory. It arises naturally by requiring that it be orthogonal to the permanent shocks[8]. Hence, the transitory shock is given by:

$$\tau_t = \frac{\boldsymbol{\alpha}'\Omega^{-1}}{\sqrt{\boldsymbol{\alpha}'\Omega^{-1}\boldsymbol{\alpha}}}\boldsymbol{\varepsilon}_t, \tag{22}$$

where the denominator ensures that $var(\tau_t) = 1$.

The two approaches to the identification of equilibrium real exchange rates that we have discussed in this section differ in one particularly important respect: the econometric permanent component we identify from an interdependent system, such as a cointegrated VAR, implicitly takes account of the fact that misalignments may feed back into economic fundamentals. As a result, the permanent component of the real exchange rate that gets identified from a multivariate decomposition may not represent the equilibrium rate that would prevail if certain shocks that are considered 'non-fundamental' by the econometrician had not occurred. Conversely, this is exactly what the CG decomposition does. It does not ask what the typical path of the system, back to equilibrium would be. But rather, it asks what the state of the system would be, if certain 'non-fundamental' shocks had not occurred.

In this paper we will employ the PT and the CG approaches simultaneously. Comparing these concepts of misalignment will then enable us to say something about the extent of hysteresis that is induced by disturbances to the real exchange rate: if misalignments are small under the cointegration-based measure, but large under the Clarida-Gali measure, then shocks that drive the exchange rate away from its fundamental value (in the Clarida-Gali concept) have an impact on the permanent value of fundamentals. Furthermore, we can then compare the persistence of the exchange rate misalignments that is implied by the various concepts.

5 The Data and their properties

We use quarterly data for the G7 countries, the United States, Japan, Germany, France, Italy, the United Kingdom and Canada, over the period 1978:Q2 to 1997:Q4. The real exchange rate series are real effective exchange rates from the IMF's International Financial Statistics (line reu), the output data measure real GDP and are also from International Financial Statistics, denominated in domestic currency (code 99B).

[8] The long-run response of the system to the reduced from disturbance $\boldsymbol{\varepsilon}_t$ is given by $\mathbf{C}(\mathbf{1}) = \mathbf{A}_0\boldsymbol{\alpha}'_\perp$ which gave us $\boldsymbol{\pi}_t = \boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_t$.The transitory shock τ_t has to be in the null space of $\mathbf{C}(\mathbf{1})$, i.e. $\mathbf{C}(\mathbf{1})\tau_t = 0$. This is the case for $\tau_t = \boldsymbol{\alpha}'\Omega^{-1}\boldsymbol{\varepsilon}_t$ because $\mathbf{E}(\boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_t\tau_t) = \mathbf{E}(\boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}'_t\Omega^{-1}\boldsymbol{\alpha}) = \mathbf{E}(\boldsymbol{\alpha}'_\perp\Omega\Omega^{-1}\boldsymbol{\alpha}) = 0$.

The construction of an output differential vis-a-vis the rest of the world would require that we use some measure of real exchange rates to construct the RoW aggregate. This might induce some spurious comovement between the real exchange rate and the output differential that we want to avoid. We therefore constructed the real output differential as the cumulated growth differential vis-a-vis the United States. For the United States, we looked at the cumulated growth differential vis-a-vis Japan and a weighted Japan-Germany average. The results were almost identical. The results for the United States reported in this paper are based on the differential vis-a-vis Japan. We do not deflate the individual country GDP's by population and so we use actual, rather than per capita, real GDP. Given that the population numbers are relatively constant for the countries and sample period used in this paper this does not affect any of the results reported below.[9]

The nominal interest rates are long bond yields (line 61) and the price indices are consumer prices (line 64). Foreign prices and foreign interest rates are in 'effective' units and have been constructed by aggregating the remaining G7 countries (i.e. exclusive of the home country) using the weights implicit in the effective exchange rates. Furthermore, we used the consumer price indices and long-term real interest rates to construct real interest rates: for each country we estimated a VAR in the process $Z_t' = \left[\begin{array}{cccc} i_t, & i_t^*, & \Delta p_t, & \Delta p_t^* \end{array}\right]$ and approximated $E(\Delta p_t)$ through forecasts from this VAR.

In specifying the appropriate lag length of the VAR in $\mathbf{X}_t$, we relied on standard information criteria. Since all of those suggested the use of either 2 or 3 lags for all countries, we decided to estimate the VAR with 2 lags throughout and to include a set of seasonal dummies. For Germany we also included a step dummy starting in 1990:Q1 to account for the effects of German reunification. In the VAR for Italy, we included a step dummy after 1992:6 to take account of the effect of the EMS-crisis.

Using a VAR specification with an unrestricted constant and without trend, we then proceeded to implement Johansen's test for cointegration. Table 1 contains the results. The data support the presence of at least one cointegrating relationship for five countries, with Japan and Canada being exceptions. We decided to follow the direction that is pointed out by our theoretical model and we imposed one cointegrating relationship throughout.

[9] The results using per capita real GDP differentials are available from the authors on request.

6 Empirical Results

6.1 Cointegrating Relationships and Variance Decompositions

As mentioned earlier, we imposed one cointegrating relationship in the estimation of all seven models. In light of the discussion in section 3, it is of particular interest to check whether i) the real interest differential is I(0) and ii) whether there is a genuine cointegrating relation between the real exchange rate and the real interest rate differential, i.e. y can be excluded from the cointegrating relationship. The results of these tests are given in table 2.

It is interesting to note that for none of the countries can we actually reject the non-stationarity of the real interest rate differential, i.e. the joint exclusion restriction $\beta_1 = \beta_2 = 0$ is strongly rejected, thereby implying that the real interest differential is non-stationary. This finding is in stark contrast to the literature, discussed in Section 2, which has, in general, failed to establish the link suggested by sticky price theories of exchange rate determination (in particular, Dornbusch's (1976) overshooting model). However, the way in which the real interest rate enters the equilibrium relation is different across countries. For three countries - the US, France and the UK - we cannot reject the exclusion restriction on y, implying that the data support the presence of a genuine cointegrating relationship between q and $r - r^*$. For the other countries in our cross-section the relationship between real interest rates and real exchange rates cannot be adequately captured without accounting for the real output differential.

Even though table 2 suggests that the data allow us to restrict the model further in individual cases, we actually left the estimated cointegrating vector unrestricted as we moved on to identify the model.

In figure 1 we provide the 'typical', i.e. cross-sectionally averaged dynamic response of the system for all G7 economies. The two permanent shocks pass the 'duck test'[10]: the real shock generally leads to a permanent increase in output relative to the rest of the world, coupled with a real appreciation (note that since we are using real effective exchange rates here an appreciation corresponds to an increase in the exchange rate) and an increase in real return. This is in line with what one should see in response to a shock to, say, total factor productivity. As for the nominal shock, this produces a temporary expansion in output coupled with an impact depreciation and a decrease of the real interest rate, and this is in line with what one would

[10] If it walks like a duck and quacks like a duck, it might actually be a duck.

expect from a monetary disturbance.

For the temporary shock a structural interpretation is not very straightforward: this shock will temporarily widen the interest rate differential and lead to a depreciation coupled with a temporary increase in output. This could possibly be read as the response to a positive fiscal shock.

However, given the aggregated nature of our model, we do not interpret the identified shocks as technology, money and fiscal shocks. Rather, we think of them as amalgams and we refer to them as real, nominal and transitory shocks throughout the remainder of the paper. What is interesting, however, is that both real and nominal shocks can have potentially lasting effects on the exchange rate.

As we move on to the variance decompositions of ΔX_t, contained in table 3, a few interesting points stand out: the second permanent shock - the nominal or monetary disturbance - explains quite a sizeable proportion of real exchange rate variance at all forecast horizons and for most of the seven countries. This finding is in line with Rogers (1999) who finds that monetary shocks - as one prime representative of nominal shocks - are an important source of real exchange rate variability. Economic theory emphasizes the role of monetary shocks for real exchange rate dynamics, but much of the earlier work, most notably Clarida and Gali (1994), could not empirically establish this important result. In US data, Rogers finds that monetary shocks explain between 20 and 60 percent of exchange rate variability and fiscal and productivity shocks combined account for between 5 and 25 percent. Our results confirm these findings for a cross-section of seven economies.

It is noteworthy how stable the share of variance that is explained by each permanent shock is over time. If there is any variation over time, it takes place in the first four quarters after which the variance shares of the various shocks reaches its permanent level.

6.2 Statistical properties of misalignment measures

Figures 2-8 give the results of the various real exchange rate decompositions for the G7. Panel a) plots the real exchange rate against the permanent component extracted using (19), i.e. the result obtained using the PT approach. Panel b) contains the corresponding results for the CG decomposition. Panel c) plots the two measures of misalignment, i.e.$q-PT$ and $q-CG$ against each other. Following Clarida and Gali, in the computation of CG we are treating all shocks except the real shock as non-fundamental. The visual impression gained from these figures is that for four of the countries (namely, France, Japan, the US, and the UK) the two measures of misalignment are very similar, although not exactly the same. For the remaining three countries

(Canada, Germany and Italy) there would appear to be important divergences in the two measures of misalignment for particular periods.

In table 4 a) and b) we give important descriptive statistics for the two measures of misalignment: the mean of their absolute values, their variance, autocorrelation and cross-correlation, as well as the half-life implied by this autocorrelation. The two misalignment measures are generally highly, although far from perfectly, correlated. The CG measure is generally more persistent. The difference in half-life is particularly pronounced for Japan, Germany, Italy, the UK and Canada. Hence, the data suggest that non-fundamental economic shocks have a pronounced effect on fundamentals and are therefore likely to change the equilibrium exchange rate.

The results in table 4 would seem to go a considerable way towards a resolution of the real exchange rate puzzle; i.e. the perceived slow mean-reversion of real exchange rates referred to in Section 2. In our cross section of G7 economies the average half-life of $\boldsymbol{\beta}'\mathbf{X}_t$ is no more than 6 quarters. We find a half-life of the CG-type transitory component close to 11 quarters, on average. In particular the numbers arising from the PT-decomposition are considerably lower than those reported in empirical studies which extend the span of the data, either by taking a long historical run of data or by using panel methods. Our results should therefore be quite relevant to the discussion on the real exchange rate puzzle: against the background of an undoubtedly very parsimonious representation of macroeconomic fundamentals, the persistence of real exchange rates appears far from excessive.

By using a forecast error variance decomposition of changes in misalignment of $\Delta\boldsymbol{\beta}'\mathbf{X}_t$, we can shed some light on the sources of shocks driving the observed equilibrium error, $\boldsymbol{\beta}'\mathbf{X}_t$. This decomposition is presented in table 5. It would appear that transitory shocks account for the bulk of the forecast variance in misalignment *changes*. In some cases, such as France and Canada, nominal shocks seem to play a role as well. Only in the UK does the real shock contribute in a meaningful way to the equilibrium error variance. Very much as in the case of Δq, the variance shares of shocks are very stable across forecast horizons. The average contribution of permanent shocks to the overall variance of misalignments mostly ranges from around forty to seventy percent.

A very interesting picture emerges from our results: the comparison of the structural CG-type misalignment measure with that obtained using only statistical information for identification, reveals that indeed only real shocks drive the non-stationarity of the data, including the real exchange rate. On the other hand, our findings resuscitate the Clarida-Gali result of an important role of nominal shocks for the variability of real exchange rates (i.e. the variance of their changes). We conclude that nominal shocks are the main determinant of high frequency movements in the real exchange rate but generally do not matter for its level or some notion of lasting disequilibrium. In spite of a very parsimonious representation of fundamentals, our measures of misalignment are considerably less persistent than the ones suggested by studies that take purchasing power parity as the reference equilibrium exchange rate.

7 Conclusion

In this paper we have examined the interaction of G7 real exchange rates with real output and interest rate differentials in the context of a structured VAR. Even though simple macro models would suggest these two variables as key determinants of real exchange rates, the literature has so far not examined them jointly in one compact econometric framework. A novel feature of our approach is that we exploit cointegration between real differentials and the real exchange rate to identify transitory and permanent components of the real exchange rate. In our analysis we generally find support for one cointegrating relationship between the output and real interest rate differentials and the real exchange rate. In some countries, this cointegrating relationship can be restricted to the real exchange rate and the real interest differential alone. We believe that this is an interesting finding since much of the earlier literature (see e.g. Baxter (1994)) could not establish this link in a bi-variate context.

Cointegration enables us to identify equilibrium exchange rates as the permanent component of real exchange rates that is consistent with dynamic equilibrium. We compare our approach to the construction of an equilibrium exchange rate to that pioneered by Clarida and Gali (1994). We find the two notions to be largely in line with each other. Our results corroborate the finding of Clarida and Gali that nominal shocks matter a lot for real exchange rate fluctuations, but we only support this claim for high-frequency movements. Using our more general framework that also allows for non-stationarity, we demonstrate that nominal shocks do not matter for the level of the real exchange rate or any notion of a persistent misalignment. In fact, in spite of the parsimonious representation of fundamentals chosen in our model, we find misalignments to be much less persistent than earlier studies. We believe these results go a large way towards explaining the so-called PPP puzzle. Furthermore, we find that it is real shocks which determine the level of the real exchange rate and this finding would seem to have an important bearing on the debate about the sources of persistence in real exchange rates (see Stockman (1987)).

Our results demonstrate that treating the real interest rate differential as an integrated variable can be a useful empirical strategy. Standard sticky price models will generally not be able to rationalize this non-stationarity but we have argued that slowly changing stances of monetary policy and financial market disturbances can make the real interest rate observationally equivalent to an integrated process in typical macroeconomic sample sizes. In this paper, we have turned this apparent problem into a virtue by exploiting it for the identification of a compact econometric system.

Summing up, it seems that real differentials provide a parsimonious representation of fundamentals for real exchange rates. Obviously, being parsimonious forbids us to assign a very specific structural interpretation to the various shocks we identify. In particular, one way our work could be extended in the future would be to explicitly recognize the separate roles of monetary and fiscal policy shocks.

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Table 1: Tests for Cointegration

h	US	Japan	Germany	France	Italy	UK	Canada	Critical Value	
				Trace Test				90%	95%
1	30.63*	26.12	59.21**	36.94**	32.10**	44.19**	25.95	28.43	31.25
2	15.14	11.63	13.19	15.49	15.79*	16.67*	7.10	15.58	17.84
3	1.61	2.18	4.13	5.81	6.12	6.60	1.70	6.69	8.08
				Maximum Eigenvalue Test					
1	15.48	14.48	46.01**	21.45**	16.30	27.5165**	18.85	18.96	21.28
2	13.54	9.46	9.07	9.67	9.68	10.0683	5.40	12.78	14.60
3	1.61	2.18	4.13	5.81	6.12	6.60	1.70	6.69	8.08

h denotes the number of cointegrating relationships

Table 2

Tests of exclusion restrictions on $\beta' = [\ \beta_1,\ \ \beta_2,\ \ \beta_3\ ]$

	$y : \beta_1 = 0$	$y\&q : \beta_1 = \beta_2 = 0$
United States	**0.16**	0.01
Japan	0.03	0.05
Germany	0.00	*0.00*
France	**0.62**	*0.00*
Italy	0.03	0.04
UK	**0.51**	*0.06*
Canada	0.00	0.00

Values reported are p-values. Accepted restrictions in bold.

Table 3
Variance decomposition of Δq_t

share of permanent shocks in forecast error					
		1	4	10	20
US	real	0.36	0.45	0.44	0.44
	nominal	0.55	0.46	0.46	0.46
Japan	real	0.25	0.26	0.24	0.24
	nominal	0.26	0.23	0.22	0.22
Germany	real	0.54	0.45	0.45	0.45
	nominal	0.43	0.51	0.52	0.53
France	real	0.01	0.09	0.15	0.15
	nominal	0.0	0.09	0.13	0.13
Italy	real	0.0	0.06	0.07	0.07
	nominal	0.73	0.70	0.69	0.69
UK	real	0.44	0.44	0.45	0.45
	nominal	0.10	0.16	0.16	0.16
Canada	real	0.02	0.04	0.04	0.04
	nominal	0.32	0.36	0.40	0.40

Table 4:
Misalignments in the G7 - descriptive statistics

a) PT measure	Mean . Abs. Value	Variance	Auto- Correlation		Half-Life (Quarters)	Cross Corr. w. CG
US	0.04	0.0029	0.90	(0.05)	6.5	0.85
Japan	0.10	0.0148	0.90	(0.05)	6.7	0.72
Germany	0.02	0.0004	0.89	(0.05)	6.3	0.12
France	0.04	0.0028	0.92	(0.05)	8.7	0.83
Italy	0.026	0.0010	0.92	(0.05)	8.5	0.34
United Kingdom	0.070	0.0093	0.78	(0.07)	2.9	0.48
Canada	0.047	0.0029	0.85	(0.06)	4.17	-0.09
			avg. half life		6.3	

b) CG-measure	Mean Abs. Value	Variance	Auto- Correlation		Half Life (quarters	Cross Corr. w. PT
US	0.03	0.002	0.83	(0.07)	4.6	0.85
Japan	0.08	0.011	0.92	(0.05)	10.3	0.72
Germany	0.05	0.004	0.95	(0.04)	10.8	0.12
France	0.03	0.001	0.90	(0.06)	8.7	0.83
Italy	0.086	0.01	0.97	(0.03)	20.5	0.34
United Kingdom	0.065	0.006	0.91	(0.05)	7.6	0.48
Canada	0.067	0.0068	0.99	(0.03)	181	-0.09
			avg. half life: (excluding CN)		10.4	

Table 5
Variance decomposition of $\Delta\boldsymbol{\beta}'\mathbf{X}_t$

share of permanent shocks in forecast error in %					
		1	4	10	20
US	real	0.03	0.09	0.10	0.10
	nominal	0.11	0.12	0.11	0.11
Japan	real	0.07	0.09	0.09	0.09
	nominal	0.01	0.02	0.01	0.01
Germany	real	0.03	0.03	0.03	0.03
	nominal	0.11	0.11	0.11	0.11
France	real	0.16	0.25	0.27	0.27
	nominal	0.26	0.28	0.28	0.28
Italy	real	0.44	0.44	0.45	0.45
	nominal	0.03	0.03	0.03	0.03
UK	real	0.49	0.48	0.48	0.48
	nominal	0.30	0.32	0.31	0.31
Canada	real	0.13	0.15	0.14	0.14
	nominal	0.50	0.49	0.50	0.50

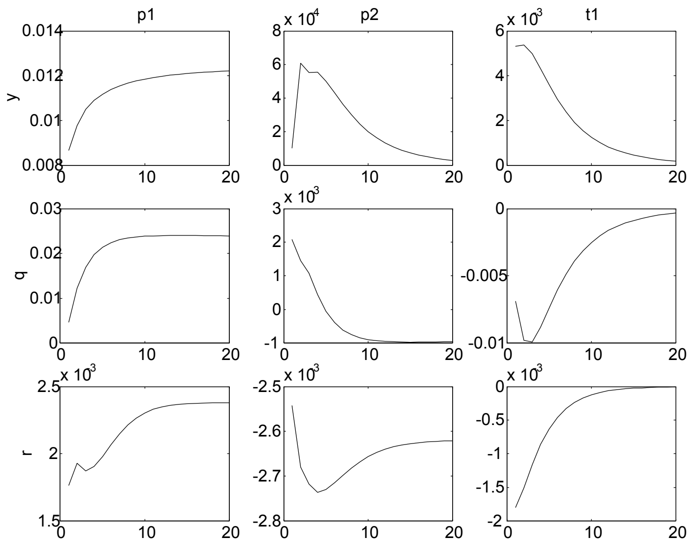

Figure 1: Cross-sectionally averaged impulse responses for the G7 economies.

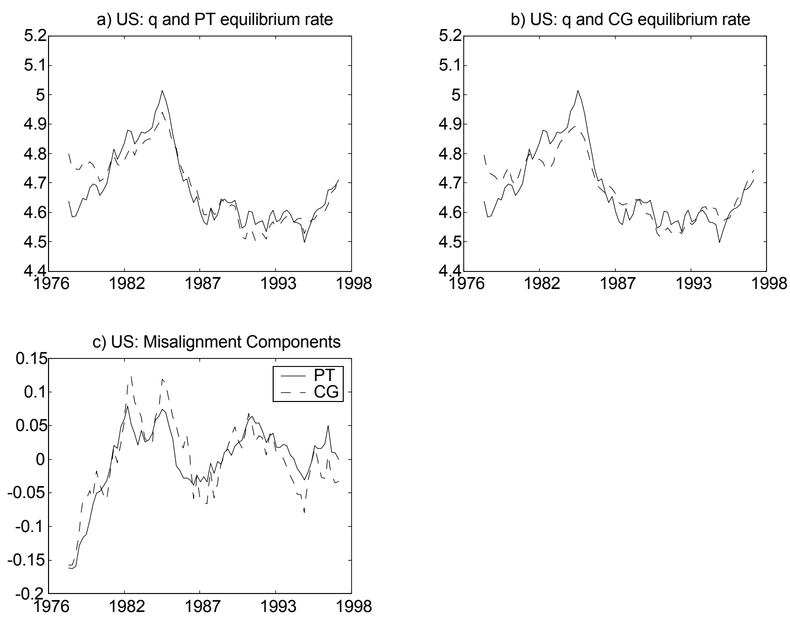

Figure 2: Decomposition of log effective real exchange rate - United States

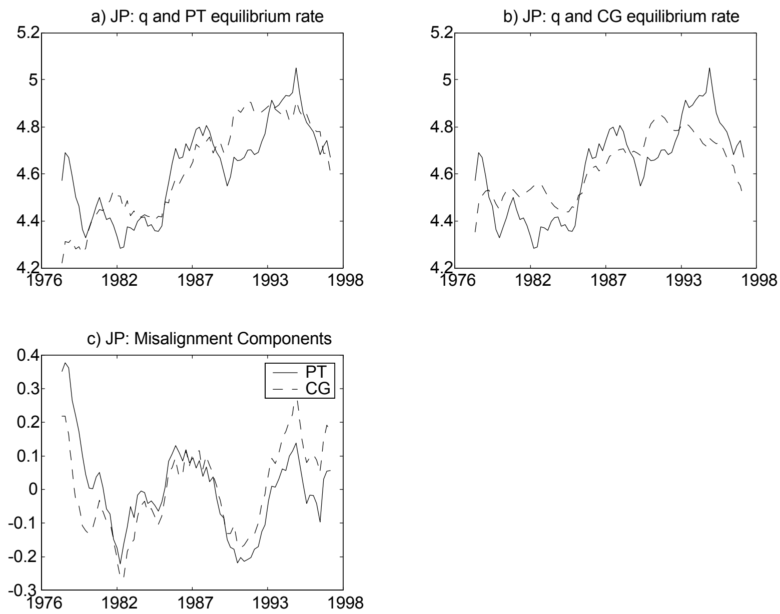

Figure 3: Decomposition of log effective real exchange rate - Japan

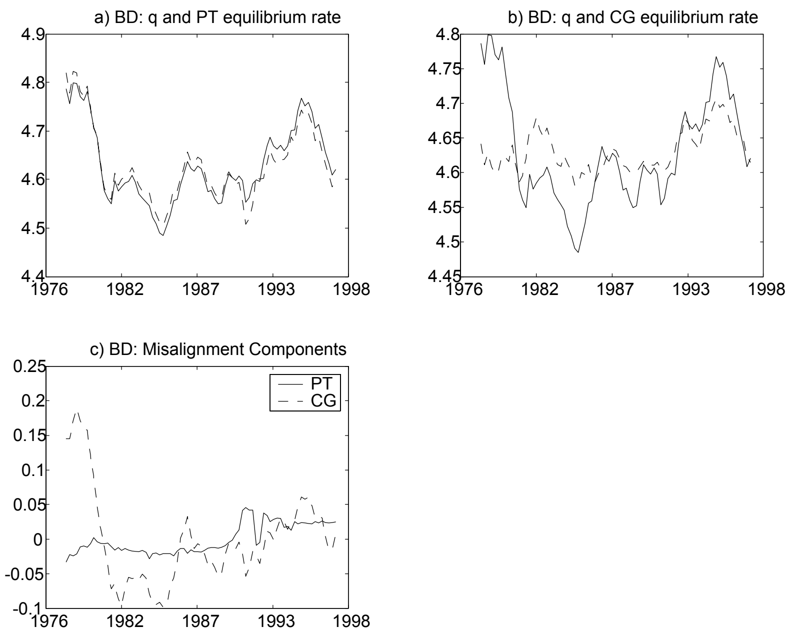

Figure 4: Decomposition of log effective real exchange rate - Germany

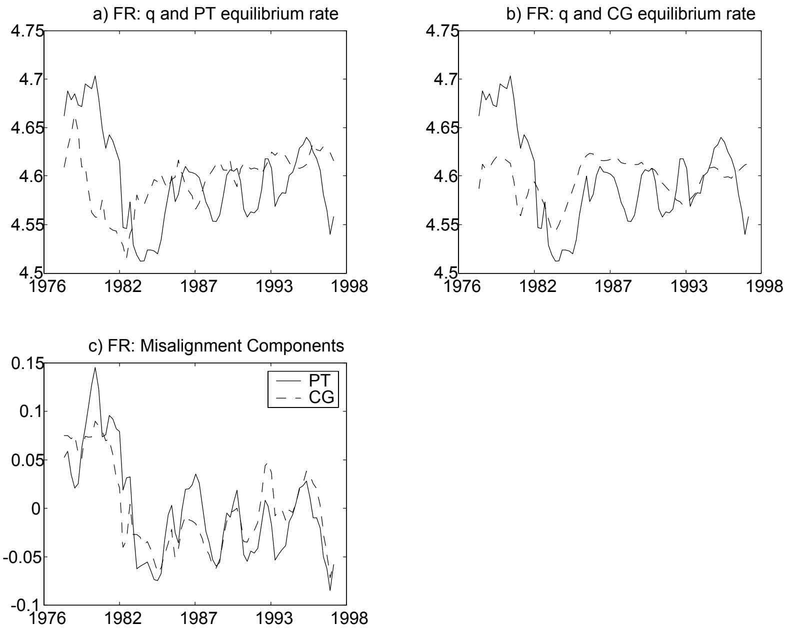

Figure 5: Decomposition of log effective real exchange rate - France

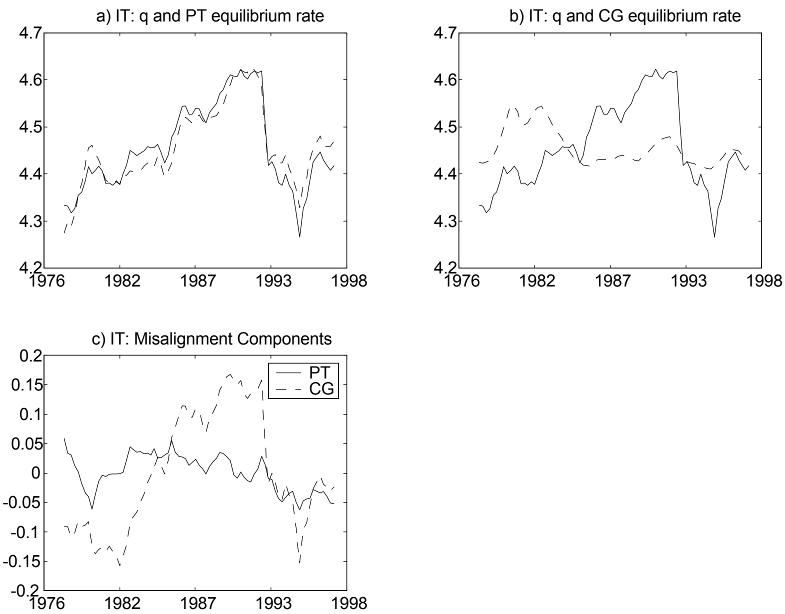

Figure 6: Decomposition of log effective real exchange rate - Italy

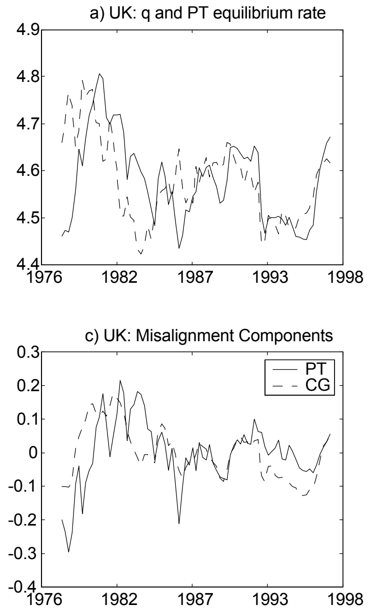

Figure 7: Decomposition of log effective real exchange rate - United Kingdom

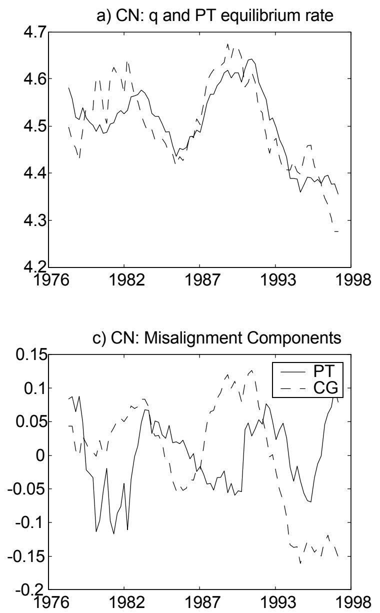

Figure 8: Decomposition of log effective real exchange rate - Canada