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THE RISE OF THE YEN VIS-À-VIS THE ("SYNTHETIC") EURO: IS IT SUPPORTED BY ECONOMIC FUNDAMENTALS?

BY CHIARA OSBAT, RASMUS RÜFFER AND BERND SCHNATZ

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Abstract: *This paper examines the long-run determinants of the euro-yen exchange rate. Using cointegration analysis, we find a consistent and significant relationship between the real exchange rate and relative productivity, the net foreign asset position, relative government spending and terms of trade shocks, as well as a fairly rapid mean reversion of the exchange rate to its equilibrium. The "equilibrium" rate tracks the trends in the actual exchange rate quite well, accounting for a large part of the yen appreciation from the mid-1970s to 2001. Our findings suggest that the euro appreciation against the yen in 2001 represented an equilibrium correction of its previous depreciation. Moreover, the width of the error bands highlights the difficulties arising when attempting to determine the precise equilibrium value of a currency.*

Keywords: *Yen, Euro, equilibrium exchange rate, BEER, cointegration.*

JEL: F31, C32

Non-technical summary

This paper examines the long-run determinants of the euro-yen exchange rate, thereby complementing the existing literature by an explicit study of the "cross-rate" within the G3 currency triangle, which does not involve the US-dollar. A direct model of this "cross rate" is also useful for understanding the depreciation of the euro vis-à-vis the Japanese yen between January 1999 and September 2000, which, in fact, has been even more pronounced than that against the US-dollar.

The paper presents a stylised model of the various structural relationships between the real exchange rate and macroeconomic fundamentals. According to this model, an increase in the net foreign asset position of a country relative to that of the other country or productivity gains relative to the partner country should lead to an appreciation of its currency. Furthermore, an increase in real interest rates should result in an appreciation. The effects of the real price of oil and of relative government spending are *a priori* ambiguous. The effect of an oil price increase depends on the relative oil dependence of the two countries and should lead to an appreciation of the less dependent country's currency. As to the effect of government spending, in the short to medium run a spending increase could support the currency through the increased demand for non-traded goods, while in the longer run confidence effects may result in a depreciation of the currency.

The empirical analysis is based on quarterly data covering the period 1975Q1 to 2001Q4. For the period prior to the introduction of the euro, a synthetic euro is constructed by using a trade-weighted average of the various legacy currencies. In addition, the model is estimated for the D-Mark-yen exchange rate in order to assess the robustness of the results.

Using cointegration techniques to estimate a so-called Behavioural Equilibrium Exchange Rate (BEER) for the euro-yen currency pair, the analysis finds a robust relationship between the real exchange rate and various economic fundamentals. In particular, the estimation results confirm the importance of relative productivity developments, the net foreign asset position and government spending (as a share of GDP) as important determinants of the real euro-yen exchange rate. In addition, the real price of oil appears to be an important explanatory variable, while the interest rate differential fails to qualify as an important factor for the real euro-yen exchange rate. The reversal of the actual exchange rate to its equilibrium value is fairly rapid, with half of any deviation from equilibrium being eliminated within less than one year. The robustness of the results notwithstanding, the size of the error bands around the estimated equilibrium exchange rate illustrates the uncertainty surrounding the estimates and suggests a cautious interpretation of the results.

The estimated equilibrium euro-yen exchange rate tracks the development of the actual exchange rate quite well between the mid-1970s and the end of the 1990s, while the relationship between the two rates becomes rather peculiar toward the end of the sample period. In 1999-2000, rising oil prices contributed to the downward pressure on the equilibrium real exchange rate of the yen vis-à-vis the euro, only partly offset by the higher Japanese net foreign asset position resulting from high and rising current account surpluses. As a result, an increasing discrepancy between the actual exchange rate and the estimated equilibrium rate materialised in 2000, and the subsequent depreciation of the yen in 2001 has constituted mainly a correction of its previous deviation from equilibrium.

For the sample period as a whole, the trend rise of the yen appears to be driven by a number of different variables, with the traditional productivity explanation being an important factor but by no means the only one. The estimates suggest that the substantial real appreciation of the yen – especially pronounced during the 1980s – was also driven by the accumulation of foreign assets and developments in oil prices. In the 1990s, the appreciation of the equilibrium exchange rate moderated, reflecting a weakening of productivity developments in Japan relative to those in the euro area, associated at least partly to the burst of the asset price bubble in Japan in the early 1990s.

1 Introduction

The depreciation of the euro after its launch in January 1999 triggered a renaissance of academic studies on the determinants of exchange rates. Most of this euro-focused literature has computed so-called equilibrium exchange rates vis-à-vis the US dollar or based on the effective exchange rate of the euro.[1] As these studies assess exchange rate developments from an equilibrium point of view, they necessarily have to take a longer-term perspective based on long-term fundamental exchange rate determinants. For the euro exchange rate, this obviously means that a "synthetic" surrogate has to be employed for the period before 1999.

Unlike the euro-US-dollar and the effective euro exchange rate, the analysis of the exchange rate of the yen vis-à-vis the euro or its legacy currencies has attracted surprisingly little academic interest in the past. There have been several studies on the bilateral equilibrium exchange rate of the yen against the US dollar,[2] but there is to our knowledge no empirical analysis available for assessing the euro-yen exchange rate.[3] Such a framework would be warranted, however, if one wants to assess, for instance, the depreciation of the euro against the Japanese currency between January 1999 and September 2000, which in fact was even steeper than against the US dollar. Moreover, given the tri-polarity in foreign exchange markets – with the US dollar, the euro and the Japanese yen as the major world currencies – the lack of a systematic analysis of the yen-euro exchange rate constitutes an important gap; a gap which this paper attempts to fill by focusing *exclusively* on this currency pair.

The paper is organised as follows: A stylised model is presented in section 2, which suggests that the real exchange rate is a function of various economic fundamentals. A reduced-form specification is subsequently presented and used as the basis for the empirical analysis. In section 3, a VECM is estimated in the tradition of the "Behavioural Equilibrium Exchange Rates (BEER)" modelling framework advocated by Clark and MacDonald (2000). The estimation results confirm the importance of relative productivity developments, the net foreign asset position and government spending (as a share of GDP) for the real euro-yen exchange rate, while the real interest rate differential fails to qualify as a determinant of the real euro-yen exchange rate. In addition, the real price of oil appears to be an important factor for the real euro-yen exchange rate. The reversion of the exchange rate to its equilibrium is

1 See Box 2 in ECB (2002) for an overview.

2 A study by Kong (2000) analyses the D-mark/yen exchange rate from a PPP perspective. On the yen-dollar exchange rate see, for instance, Yoshikawa (1990), Chinn (1997), MacDonald and Nagayasu (1998), Kasuya and Takagawa (2001) and De Carvalho (2002). On the effective yen exchange rate, see Nagayasu (1998).

3 In fact, there are only few studies which build on more comprehensive projects and allow to derive implicitly equilibrium exchange rates for the yen against the euro. Driver and Wren-Lewis (1998), for instance, use a globally consistent framework based on "Fundamental Equilibrium Exchange Rates" to compute a FEER for the yen against the D-mark. Alberola et al. (1999) conduct a panel study based on so-called "Permanent Equilibrium Exchange Rates (PEER)" and assess the exchange rate level in 1998Q4.

fairly rapid, with a half-life of less than one year. In section 4 actual exchange rate developments are compared with the estimated BEER and its confidence bands, which are constructed using a method recently proposed by Johansen (2001) and which is theoretically consistent with the VECM framework. We find that the (equilibrium) trend appreciation of the yen moderated in the 1990s as a result of weak productivity developments in Japan, associated at least partially to the burst of the asset price bubble in Japan in the early 1990s. In 1999/2000, rising oil prices have also contributed to the downward pressure on the equilibrium real exchange rate of the yen vis-à-vis the euro, only partly offset by the higher Japanese net foreign asset position resulting from high and rising current account surpluses. As a result, an increasing discrepancy between the actual exchange rate and the BEER materialised in 2000, and the subsequent depreciation of the yen in 2001 has constituted mainly a correction of its previous appreciation.

2 A model for the real exchange rate

2.1. *Balance of payments equilibrium and the real exchange rate*

A prominent channel through which fundamentals may influence trends in the real exchange rate is related to macroeconomic balance considerations, as implied in portfolio balance models proposed by Branson (1977), discussed by Frenkel and Mussa (1985) and applied by Faruqee (1995), Fell (1996) and MacDonald (1999). In these models, equilibrium in the balance of payments is a central determinant of the real exchange rate, which moves to equilibrate the individual items in the balance of payments (current and financial account) to ensure stock and flow market equilibrium. In general, the change in the net foreign investment position is equal to the current account balance, which in turn consists of the trade balance and the balance in the income account.[4]

(1) $$\Delta F_t = ca_t = nx_t + r^* F_t,$$

where ΔF_t is the change in the net foreign asset position, ca_t the current account position and nx_t represents the balance of trade. F_t is the outstanding stock of net foreign assets which – on average – is assumed to yield an interest rate of r^*, the given world interest rate. The balance of trade is determined as

(2) $$nx_t = -\lambda q_t + \mu c_t$$

where c_t is a shift factor, incorporating exogenous variables affecting foreign demand for domestic goods and domestic demand for foreign goods, q_t is the real exchange rate, and λ, μ

4 Current transfers are, in general, comparatively small and of limited economic content. For modelling purposes, they are therefore excluded in the discussion of the current account.

represent the sensitivity of net exports to movements in these variables.[5] Following Faruqee (1995), this shift factor may include variables such as productivity growth differentials in the non-traded and traded goods sectors or commodity price shocks affecting the terms of trade. Other possible exogenous factors may relate to distortions arising from unbalanced government spending, which may give rise to expectations of distortionary taxes in the future. Net exports are negatively affected by the real exchange rate (defined as $q_t \equiv s_t + p_t - p_t^*$ where s_t is the nominal exchange rate expressed in foreign currency per unit of domestic currency and p_t and p_t^* are the domestic and the foreign price index, respectively).

The financial account position, which is the counterpart of the current account balance, is assumed to be a function of the discrepancy between the target level of net foreign assets, $\overline{F}$, and their current actual level, F, and of the expected change in the real exchange rate. The target level is assumed to be exogenous and fixed in the current setting but it could also be linked to factors such as the demographic profile of a country. More formally, the financial account (fa_t) can be written as:

(3) $$fa_t = \kappa(\overline{F}_t - F_t) + \gamma E_t(\Delta q_t)$$

where κ represents the speed of adjustment from the actual to the desired level of net foreign assets.

Assuming flexible exchange rates, which ensures that the change in foreign exchange reserves equals zero, the current and financial account balance must match. Accordingly, equations (1) and (3) constitute a simultaneous forward-looking system of difference equations that can be solved for the expected future time paths of the endogenous variables q_t and F_t, conditional on the current inherited stock of net foreign assets and the expected future time paths of the exogenous variables $\overline{c}_t$ and $\overline{F}_t$. In the long-run equilibrium, the change in net foreign assets has reached its desired level and the trade balance equals interest payments, so that the solution for the current equilibrium real exchange rate is:

(4) $$q_t = \overline{q}_t + \eta(F_t - \overline{F}_t) \qquad \text{with } \eta > 0$$

where: $$\overline{q}_t = \frac{\mu}{\lambda}\overline{c}_t + \frac{r^*}{\lambda}\overline{F}_t .$$

Bars over the variables indicate long-run equilibrium values, while the other variables reflect current flow equilibrium values. $\overline{c}_t$ and $\overline{F}_t$ depend on the present discounted value of their future values in line with the forward-looking character of the model, which underlines the importance of future economic conditions for the determination of exchange rates. The expression in (4) suggests that the actual real exchange rate depends on the long-run

[5] In fact, the parameter λ contains the relevant price elasticities and is assumed to be positive, implicitly assuming that the Marshall-Lerner condition is satisfied. For details, see Frenkel and Mussa (1985).

equilibrium exchange rate, $\overline{q}_t$, and the divergence between the current value of net foreign assets holdings and their long-term desired level. The sustainable path for the real exchange rate may differ from its long-run value until full stock equilibrium is attained. The long-run equilibrium exchange rate itself depends on the long-run levels of the net foreign asset position and other exogenous variables.

2.2. Real interest rate differential

Following Fell (1996), equation (4) can be linked to the uncovered interest rate parity condition in real terms. Assuming that the real exchange rate reverts only gradually to its long-term equilibrium over the maturity of the bonds, the current real exchange rate can be expressed as a function of its long-term equilibrium, from which it may deviate in the medium-term owing to movements in the real interest rate differential.

(5) $$q_t = \overline{q} - (1/\theta)(r_t^* - r_t).$$

where θ represents the (expected) speed of adjustment and maturity coefficient ($0 < \theta < 1$). Without frictions in the economy, θ is, in principle, given as the maturity of the bonds. As Edison and Melick (1999) have pointed out, however, a 10-year coupon bond should be a good proxy for a roughly 7-year pure discount bond, suggesting a coefficient of roughly 7. MacDonald and Nagayasu (2000) have shown, moreover, that the parameter θ changes if there are rigidities in the economy, and the maturity effect is reduced suggesting a coefficient of less than 7 for the real interest rate differential.

2.3. Productivity developments and the real exchange rate

According to the Balassa-Samuelson theorem, productivity advances in the traded goods sector should be associated with a higher long-run real exchange rate, if the latter is based on broad price and cost indices. This effect reflects an equilibrium phenomenon, which works through the prices of non-traded goods affecting the relative internal price ratio between non-traded and traded goods q^I_t. This term derives from a decomposition of the real exchange rate into a tradable and a non-tradable component, as shown in the following.

The price level is derived from a weighted average of the prices of traded goods (p_t^T) and non-traded goods (p_t^N), where α is the share of traded goods in the domestic goods basket and α^* the share of traded goods in the foreign goods basket. Allowing for the arbitrage condition for internationally traded goods, the real exchange rate can be decomposed into domestic and foreign traded and non-traded goods prices:

(6) $$q_t = q_t^T + q_t^I = q_t^T + (1-\alpha)\cdot(p_t^N - p_t^T) - (1-\alpha^*)\cdot(p_t^{N*} - p_t^{T*}),$$

If PPP holds for traded goods, q_t^T is constant. In reality, however, it might be influenced by terms of trade shocks such as sudden changes in oil prices. The second term – q_t^I – may be influenced by diverging trends in the relative price of traded and non-traded goods in the two countries, as well as by the relative share of non-traded goods in the baskets for the price indices. This term is directly related to the Balassa-Samuelson result discussed below.

A universal starting point for deriving the Balassa-Samuelson result is the standard Cobb-Douglas production functions:

$$(7) \qquad Y_t^T = A^T (L_t^T)^\sigma (K_t^T)^{1-\sigma}, \quad Y_t^N = A^N (L_t^N)^\rho (K_t^N)^{1-\rho}$$

where σ and ρ represent labour intensity of production in the traded and non-traded sector, respectively. The model assumes that labour is mobile between sectors and is paid according to its marginal product. The first assumption suggests that wages are equalised across sectors and the second implies that the sectoral productivity differential is equal to the sectoral price differential; denoting by Z^T and Z^N average labour productivity in the traded and non-traded sector respectively, this can be written as

$$(8) \qquad \frac{\partial Y^T / \partial L^T}{\partial Y^N / \partial L^N} = \frac{\sigma Y^T / L^T}{\rho Y^N / L^N} = \frac{\sigma}{\rho} \frac{Z^T}{Z^N} = \frac{P^N}{P^T},$$

This results in the following expression in logarithms:

$$(9) \qquad p_t^N - p_t^T = \log(\sigma / \rho) + (z_t^T - z_t^N),$$

which can be substituted into (6), yielding

$$(10) \qquad \begin{aligned} q_t = \quad & q_t^T + q_t^I = q_t^T + (1-\alpha) \cdot (\log(\sigma/\rho) + (z_t^T - z_t^N)) \\ & -(1-\alpha^*) \cdot (\log(\sigma^*/\rho^*) + (z_t^{T*} - z_t^{N*})) \end{aligned}$$

Abstracting from the issue of different shares of non-tradable goods in the two countries, (10) implies that a productivity gain in the traded goods sector of a country relative to the non-traded goods sector will lead, *ceteris paribus*, to an appreciation of that country's currency. The main reason for this result is that the high productivity growth in the traded goods sector leads to a strong wage increase in both sectors. However, in the non-traded sector the wage increase is not offset by a corresponding productivity increase, pushing up the relative price of non-traded goods. Unlike in the case of price changes for traded goods, the price increase is not completely offset by a nominal depreciation of the exchange rate and the real exchange rate thus appreciates.

In sum, the individual components of the model provide guidance for the choice of variables in the empirical implementation of the theoretical framework as well as for the signs and possible magnitude of the coefficients:[6]

$$(11) \qquad q_t = \quad \beta_0 + \beta_1 aca_t + \beta_2 rid_t + \beta_3 (prod_t / tnt_t) + \beta_4 gov_t + \beta_5 tot_t,$$

6 Obviously, the reduced-form is not derived from a full structural model but rather from a pragmatic examination of potential channels through which economic fundamentals may influence the real exchange rate.

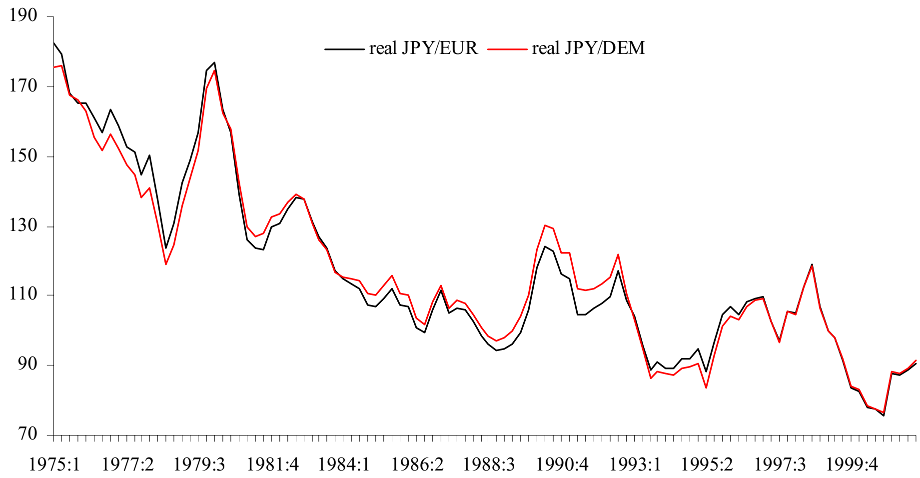
where q_t is the logarithm of the real exchange rate, aca_t stands for the relative net foreign asset position (as a ratio of GDP), rid_t is the real interest rate differential, the third term corresponds to the log of the productivity differential, measured directly as GDP over employment ($prod_t$), or, alternatively, in a more indirect way as the relative price ratio of non-traded and traded goods (tnt_t), gov_t reflects government spending (as a ratio of GDP, and in logs), and tot_t represents the (log of) terms of trade (captured in our empirical specification by oil_t, which stands for the log of the real price of oil). All variables are expressed as differentials between the euro area and Japan, except the real price of oil. β_0 is a constant term, β_1 is expected to be positive, implying that a higher net foreign asset position should be associated with a real appreciation. β_2 is expected to be negative and smaller than 7 in magnitude. β_3 should be positive and close to 1 if the economies considered are assumed to have non-traded goods sectors of similar sizes. The sign of β_4 is ambiguous, since higher government spending can have a positive effect on the real exchange rate in the short to medium run by increasing the price of non-traded goods, but should have a negative impact in the long run owing to confidence effects. The sign and the magnitude of β_5 depend on the relative oil dependence of the country under consideration.

3. Econometric Analysis

3.1. Data and measurement issues

In view of the fairly short period for which data for the euro are available, a "synthetic" euro-yen exchange rate has been computed as a weighted geometric average of yen exchange rates of the individual EMU currencies for the period before 1999. The weights for the "theoretical euro" exchange rate are based on the share of each euro area country in total manufacturing trade of the euro area vis-à-vis non-euro area countries. In Chart 1, the ("synthetic") real euro-yen exchange rate is compared with the corresponding D-mark-yen real exchange rate. The decline in the index corresponds to a gradual appreciation of the yen. The exchange rates of the D-mark and the euro vis-à-vis the yen exhibit an extremely high degree of correlation. As from the point of view of conceptual consistency over the entire sample, it appears preferable to refer to the "synthetic" yen-euro exchange rate prior to the launch of the euro rather than the D-Mark-yen exchange rate, hence we choose the former for the empirical estimation. Nonetheless, we have carried out robustness checks using the yen-D-mark real exchange rate. Estimating the same model for the yen-D-mark yields very similar coefficients for the variables that turn out to be significant. More detailed results are reported in Annex 2. The time series for the other variables referring to the euro area have been computed on the basis of the methodology described in detail in Maeso-Fernandez et al. (2001). The real exchange rate is based on consumer price indices.

Chart 1. Real yen-euro and yen-D-mark exchange rate

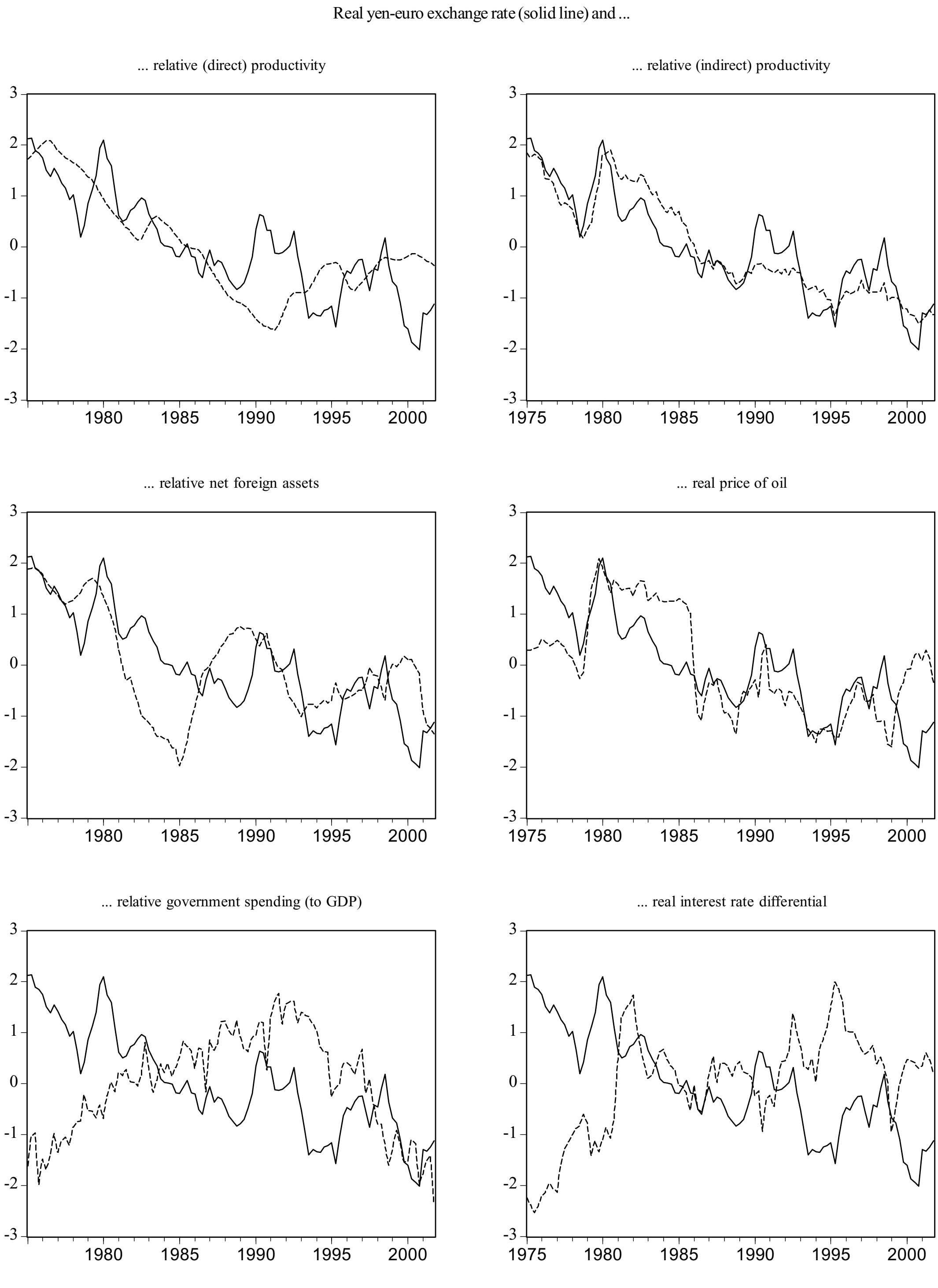


The theoretical framework presented in section 2 suggests that productivity differentials in the traded goods sector between the euro area and Japan may be a determinant of the equilibrium exchange rate, according to the Balassa-Samuelson effect. This argument should, in principle, be more relevant for exchange rates between emerging markets and industrialised countries. The yen exchange rate is a notable exception, however, since Japan experienced a significant catching-up effect in the second half of the last century, prior to the burst of the asset price bubble in the early 1990s. Measuring productivity differentials appropriately is, however, subject to statistical problems due to differences in accounting procedures as well as major conceptual difficulties. A relatively direct measure of diverging productivity trends relies on differences in total labour productivity, measured as real GDP divided by the number of employed persons. This measure has some drawbacks: first, it measures economy-wide productivity rather than productivity advances in the traded sector as implied in the Balassa-Samuelson theory.[7] Secondly, in order to take into account country-specific and time-varying preferences for part-time work, output per hour worked would be a better proxy, but limited data availability precludes this (for the time being). These conceptual difficulties notwithstanding, there is a close correlation between the real appreciation of the yen against the euro and relative total labour productivity (dotted line) up to the late 1980s (Chart 2).

[7] The assumption that total labour productivity is a good proxy for traded goods productivity in Japan may be eroded by the fact that massive public work programmes were initiated in Japan in the 1990s, which might have had a negative impact on non-traded goods productivity.

Chart 2. Real yen-euro exchange rate and economic fundamentals

(data normalised with N(0,1))



Subsequently, this correlation loosened as productivity declined in Japan in the aftermath of the downturn of the Japanese economy following the bursting of the asset price bubble. Since the launch of the euro, relative productivity appears to have stabilised, while the yen appreciated again strongly vis-à-vis the euro. A rather indirect measure of productivity differentials, which attempts to take sectoral considerations into account, is the relative price of non-traded and traded goods (Kakkar and Ogaki 1999). As outlined in section 2, the relative prices of non-traded and traded goods are inversely related to productivity developments in the two sectors. Since it is difficult to discriminate in practice between traded and non-traded goods, the consumer price index has been frequently used as a proxy for non-traded goods, whereas producer or wholesale prices have been employed as proxies for traded goods. A significant drawback of this measure is that domestic demand shocks in the non-traded goods sector, as well as changes in the tax policy, may conceal the actual productivity information conveyed by this variable. In addition, commodity price shocks may distort this variable if they have systematically a different effect on consumer and producer prices. In the multivariate analysis, therefore, we control for oil price developments. Chart 2 illustrates, however, that there has been a rather close link between the trends in this productivity measure and the real yen-euro exchange rate over the past 25 years.

The international investment position has been calculated as the cumulated current account balance in relation to GDP, and must be interpreted cautiously. First of all, euro area data have been aggregated on the basis of national data. While intra-euro area components of the individual current account position should cancel out in theory, they do not in practice. In addition, "errors and omissions", which may reflect unrecorded capital or trade flows, may worsen the mis-measurement of the international investment position. Moreover, for bilateral exchange rates, it may be argued that one should use the net foreign asset position between the two countries involved, which is unavailable. Instead, the difference between the overall net foreign asset positions relative to GDP of the euro area and Japan has been used. The simple accumulation of current account positions has the additional drawback that the impact of valuation issues and the effects of debt forgiveness and reinvested earnings on this variable are neglected.[8] The difference in the net foreign asset positions of the euro area and Japan exhibits relatively large and persistent swings. As can be seen from Chart 2, the relative improvement of the Japanese net foreign asset position between 1975 and 1985 coincided with an appreciation of the yen. There is also evidence of correlation between this variable and the real yen-euro exchange rate in the 1990s, when the appreciation of the yen was again reflected in an improvement of the relative net foreign asset position. In the second half of the 1990s, the yen depreciated in real terms as the net foreign asset position of Japan deteriorated.

Terms-of-trade shocks are captured by changes in the real price of oil, which is computed as the US-dollar oil price converted into yen and deflated by Japanese wholesale prices. A rise in this proxy variable represents a rise in production costs, reflecting the fact that oil is an important input in production. Rising oil prices, moreover, reduce the purchasing power of net oil importing countries. Since both the euro area and Japan are net oil importers, a rise in oil prices may generate an appreciation or a depreciation of the exchange rate depending on

[8] Lane and Milesi-Feretti (2001) illustrate, however, that the cumulated current account appears to provide an overall good representation of trends in the net foreign investment position.

their relative “oil price vulnerability”.[9] Overall, while the energy intensity of the Japanese economy is slightly lower than in the euro area given its commitment to energy preservation and greater reliance on nuclear power, the oil trade balance is significantly more negative for Japan than for the euro area. As a consequence, a negative effect of higher oil prices on the yen seems more likely. This is also suggested by Chart 2, which shows a rather close correlation between the real price of oil and the yen-euro real exchange rate. However, this correlation has broken down since 1997.

Following the model outlined above, this paper also evaluates whether relative government spending as a ratio of GDP (defined as euro area minus Japan) and the real interest rate differential have a significant effect on the real euro-yen exchange rate.[10] While the charts for these variables do not indicate any clear co-movement, the econometric analysis finds some evidence that relative government spending is an important variable for finding a stable long-run relationship in combination with the other variables.

3.2. Cointegration analysis and results

3.2.1. The vector-error correction model

Since the data shown in Chart 2 clearly exhibit trending behaviour and high persistence, it seems reasonable to use cointegration analysis to estimate the model. As a consequence, the maximum likelihood approach suggested by Johansen (1995) has been employed. This approach can be briefly summarised as follows. Let y_t be a vector of I(1) variables. The estimated model is a Vector Autoregression (VAR) of dimension *p* and order *k,* which can be written in vector-error-correction (VEC) form:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \mu + \varepsilon_t \,, \tag{12}$$

where Π is a matrix containing the long-run coefficients and the adjustment terms, Γ_i represent matrices of short-run coefficients, μ is a vector of constants, and ε_t denotes a vector of iid gaussian-distributed errors. If the matrix Π has reduced rank (*r*), it can be expressed as the product of a (*p x r*) matrix of loading coefficients α, and the (*p x r*) matrix β, containing the *r* cointegrating vectors, so that $\Pi=\alpha\beta'$. The matrix α gives an indication of the importance of the cointegration relationships in the individual equations of the system and of the adjustment speed, while the matrix β represents the long-term equilibrium relationships.

9 See Amano and van Norden (1999) on the relationship between the real exchange rate and the price of oil. The real price of oil was employed as a proxy for terms of trade by e.g. Clostermann and Schnatz (2000), in a study on the euro-dollar exchange rate and by Maeso-Fernandez et al. (2001) in an analysis of the euro effective exchange rate. For some quantitative analyses regarding the importance of oil in individual countries see Davies and Strongin (2000) and Jen et al. (2001).

10 The real interest rate was computed as the nominal 10-year bond yield minus the consumer price inflation rate in the previous year. The differential is defined as the euro area minus the Japanese real interest rate.

The moving average representation of the VEC model in (12) provides additional information to analyse the properties of the system. It is defined as

(13) $$y_t = C\sum_{i=1}^{t}\varepsilon_i + C\eta + C(L)(\varepsilon_t + \eta)$$

where:

(14) $$C = \beta_{\perp}(\alpha_{\perp}^{'}\Gamma\beta_{\perp})^{-1}\alpha_{\perp}^{'} = A\alpha_{\perp}^{'}$$

In these equations, $\alpha_{\perp}$ and $\beta_{\perp}$are the orthogonal complements of α and β. $\alpha_{\perp}$ spans the space of the common stochastic trends, i.e. it identifies the linear combinations of the cumulated shocks that form the common trends or driving forces of the system. The matrix $\Gamma = I - \Sigma_{i=1}^{k-1}\Gamma_i$ is a function of the short-run coefficients. The matrix A represents the loading factors of the common trends, and indicates to what extent each trend influences each variable. Finally, the C matrix measures the long-run impact of shocks to the system.

The cointegration model in (12)-(14) forms the analytical basis used to construct behavioural equilibrium exchange rates (BEER), as the fitted value from the estimated equilibrium relationship (y_{1t}-β'y_t), where y_{1t} indicates the first variable in the system, i.e. in our case the real exchange rate.

Moreover, based on Johansen (2001), error bands can be derived for the BEER on the basis of the following equation, which builds on the VEC model and its moving average representation discussed above (see annex 1 for details on the derivation):

(15) $$u'_1\hat{C}\hat{\Gamma}y_t = \pm 2\sqrt{\left((I - C\Gamma)\bar{\beta}\,\mathrm{Var}(\beta' y_t)\,\bar{\beta}'(I - C\Gamma)'\right)_{11}}$$

where y_t is the data matrix, $\bar{\beta} = \beta(\beta'\beta)^{-1}$, and C, Γ as defined above.

3.2.2. Estimation results

The VEC model was estimated over the sample running from 1975Q1 to 2001Q4. The constant has been restricted to the cointegrating space, which is equivalent to assuming that the data have no linear deterministic trend in the levels.[11] This is consistent with the descriptive statistics of the data, which suggest that some variables have non-zero means, while all the growth rates have zero means (see Table 1).[12] The time-series properties of the data have been checked within the cointegration model by having stationarity as the null hypothesis given the cointegration space. This multivariate procedure, which has the

[11] Alternatively, employing – in spite of the descriptive statistics shown in Table 1 – a model including an unrestricted constant and a restricted trend, the trace test results do not change, i.e. the tests always indicates the presence of only one cointegration relationship. Including a restricted trend in the procedure for testing for cointegration rank has the nice property that the trace test result is not contingent on the choice of deterministic specification. Furthermore, the LR test on beta clearly indicates that the restricted trend is insignificant in the cointegration relationship.

[12] Three centred seasonal dummies have been initially included to capture seasonal effects without affecting the inference on the long-run coefficients, but they turned out to be insignificant and were hence excluded from the model.

advantage of taking information from the entire data set into account, employs a LR test, which is distributed as a χ^2 with *p–r* degrees of freedom. For the selected model, the hypothesis of stationarity can be rejected within the VEC model for all variables (see Table 2).

Table 1. Descriptive Statistics, 1975Q1-2001Q4

	q	*prod*	*oil*	*aca*	*gov*	*ridl*
			Levels			
Mean	0.13626	0.01081	5.38805	-12.0234	0.73414	3.12969
Std. Dev.	0.20097	0.04383	0.48289	7.71044	0.05667	1.23738
			Differences			
Mean	-0.00601	-0.00086	-0.00285	-0.23322	-0.00039	0.02882
Std. Dev.	0.05145	0.00368	0.14447	1.40969	0.02142	0.42135

The lag order has been chosen on the basis of lag exclusion tests (F-test).[13] Accordingly, the VAR was estimated based on four lags, as the exclusion of the fourth lag is rejected at the 5% level, while the F test fails to reject the exclusion of the fifth lag. The number of cointegrating vectors is determined on the basis of the Johansen trace statistics since Cheung and Lai (1993) found that the trace test is more robust to skewness and excess kurtosis in innovations than the maximum eigenvalue test.[14]

Alternative specifications of the VEC model were initially estimated separately for each productivity measure including also all the other variables: the oil price (oil), the difference of the accumulated current account of the euro area and Japan (aca), relative government spending, and the real interest rate differential. In addition, in order to ensure stability in the relationship, it was necessary to include dummies in the short-run dynamics, which capture outliers for developments in the real price of oil ("oil dummy") and in the accumulated current account position ("aca dummy"). The entries in both dummies add up to zero, so that they do not affect the inference on the long-run coefficients or the asymptotic distribution of the cointegration trace statistics. On the real price of oil, the "oil dummy" takes the value of -1 in 1986Q1, when oil prices plunged as OPEC oil output increased sharply during a price war as OPEC failed to agree upon a production accord. The "oil dummy" also takes the value of 1 in 1990Q3, when oil prices spiked during the Iraq intervention in Kuwait. The "aca dummy" refers to periods of sudden shifts in the yen exchange rate, which had a strong and immediate impact on the Japanese international investment position. Accordingly, the "aca dummy" was set to 1 in 1998Q4, when the yen appreciated vis-à-vis the US dollar by more than 10%, its largest quarter-on-quarter depreciation in the sample period. The same dummy

[13] The results are broadly unchanged when using instead three lags in the VAR, as suggested by AIC and SIC.

[14] Using the more restrictive statistics proposed by Reinsel and Ahn (1992) and applied by Reimers (1992) in a simulation on cointegration tests, it is more difficult to find cointegration. As Doornik et al. (1999) point out, however, such a small-sample scaling of the trace test is not theoretically founded and may overcorrect the statistics in certain cases.

was set to -1 in 2001Q1, when the yen depreciated by 8% against the US currency and by 14% against the euro.

The specification including the indirect productivity variable was abandoned due to signs of misspecification, given that even the real exchange rate equation exhibits episodes of significant instability within this system. Accordingly, the following analysis focuses on the direct productivity variable, defined as GDP/employment. For this specification, the cointegration trace test based on standard critical values indicates the presence of a single cointegration vector. After imposing the rank restriction and normalising, however, the real interest rate differential was insignificant in both the α and the β vector, so that this variable was excluded from the specification, and the final model only includes five variables. Table 2 summarises the corresponding estimation results. Specification I refers to the model estimated under the restriction that the cointegration rank is one; specification II also incorporates weak exogeneity restrictions on the adjustment coefficients.[15] The choice of cointegration rank and the long-term estimates are discussed in more detail below.

Table 2: VECM results (t statistics in parentheses)

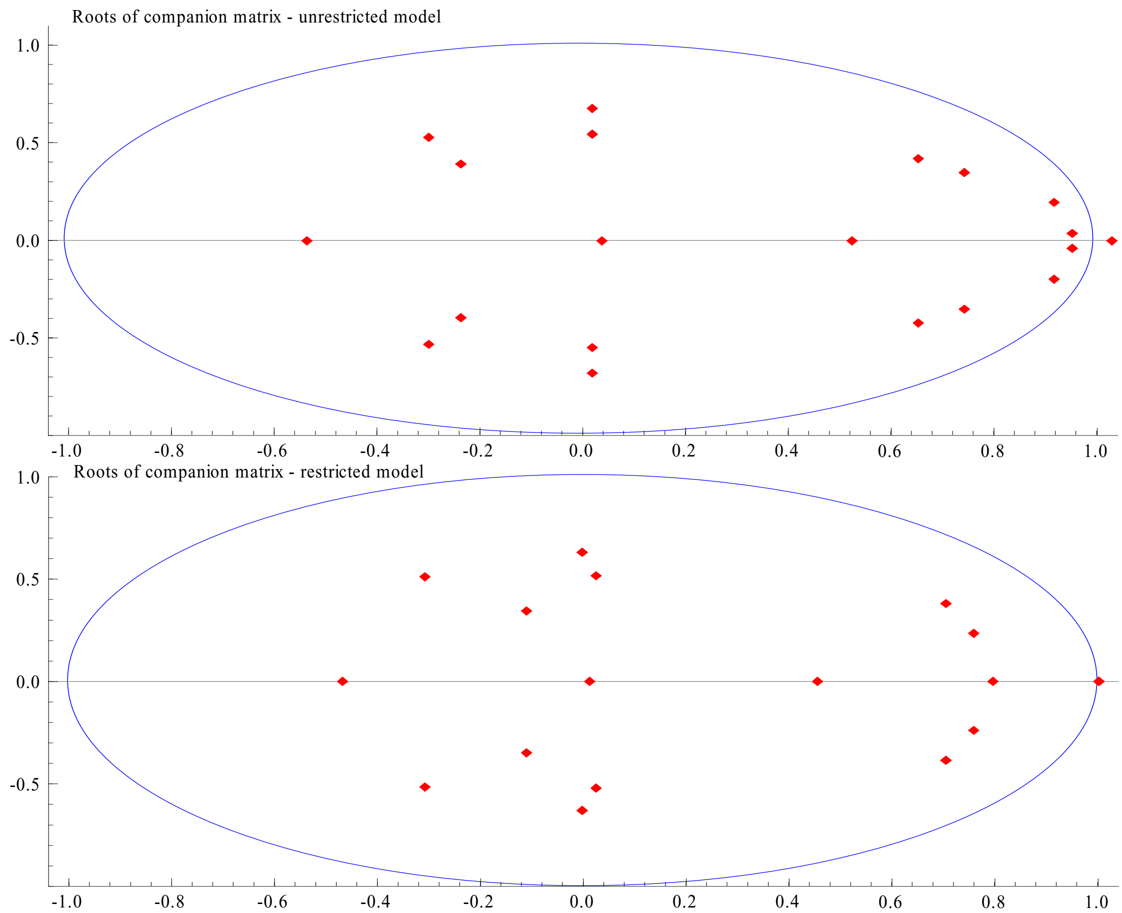
		Specification I		Specification II	
	Stationarity test	Long-term coefficients	Adjustment terms	Long-term coefficients	Adjustment terms
q	32.74	1.000	-0.153	1.00	-0.166
		-	*-2.616*	-	*-3.061*
prod	36.00	-2.901	0.005	-2.484	0.006
		-5.050	*2.973*	*-4.370*	*3.48*
oil	39.33	-0.128	-0.108	-0.164	--
		-2.981	*-0.723*	*-3.854*	--
gov	40.34	-1.354	0.072	-1.270	0.073
		-3.981	*3.281*	*-3.773*	*3.270*
aca	40.28	-0.009	-1.660	-0.011	--
		-4.480	*-1.282*	*-5.417*	--
constant		1.445		1.545	
		5.362		*5.794*	
Summary Statistics:					
Number of lags:		4		4	
Test on restriction (p-value)				0.39	
LM(1-4) (p-value)		0.06			
White test (p-value)		0.90			
Cointegration test			Standard critical values		
			5%	1%	
Trace-test	None	89.14	76.07	84.45	
	1	41.78	53.12	60.16	
	2	20.58	34.91	41.07	
	3	10.11	19.96	24.60	

Constant restricted to the cointegrating space, Critical value for stationarity test: 11.07
Standard critical values for the model with a constant restricted to the cointegrating space as suggested by Osterwald-Lenum (1992),

[15] Weak exogeneity has been identified by testing whether $\alpha = 0$ for each variable both separately and jointly. If the restriction holds, a variable is defined as weakly exogenous.

Choice of cointegration rank: The trace tests indicate the presence of one cointegration relationship at the 1% level based on standard critical values. The hypothesis of one cointegration relationship is confirmed by looking at the magnitude of the roots of the system, (Chart 3) and by the graph of the cointegration vector, which depicts $\beta' y_t$ (Chart 4).

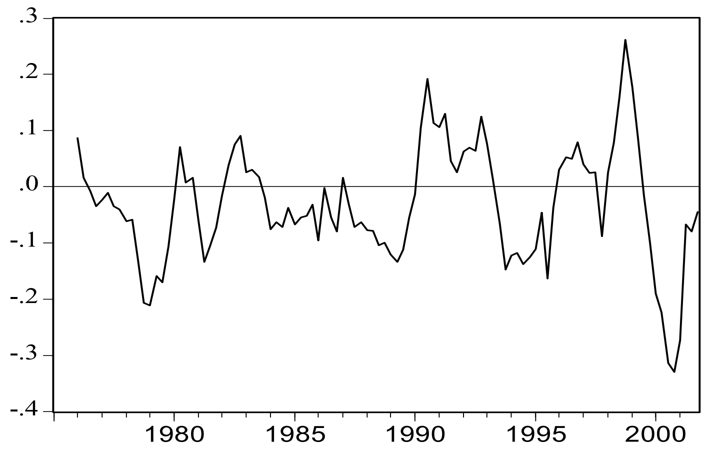
Chart 3: Companion matrix



The upper chart shows the position of the estimated roots in the unit circle for the unrestricted system, while the lower chart shows the corresponding picture after imposing the restrictions of a rank of one and two zero restrictions on α. In the unrestricted model, one root is slightly above one, and two conjugate pairs of roots have modulus larger than 0.93. In the restricted model, which implies that the system contains 4 unit roots, the fifth largest root drops to below 0.8, suggesting that the rank restriction is appropriate, as there seem to be only 4 common trends driving the system.

Chart 4 shows the cointegration vector, which appears indeed to be stationary, both by visual inspection and by standard ADF tests. This additional evidence supports the finding of a rank equal to one suggested by the standard trace test.

Chart 4: Cointegration vector ($\beta' y_t$)



Long-term relationship and adjustment: Tests for exclusion from the long-run relationship confirm that the real exchange rate cannot be excluded in any of the models. Accordingly, it is feasible to normalise its coefficient to one in the long-run relationship (see specification I in Table 2). The standard specification tests broadly suggest white-noise residuals, though there remains some autocorrelation in the errors at the 10% level. All variables are highly significant in the long-term relationship and their signs are in line with the theoretical reasoning described in section 2. *Ceteris paribus*, a 1% rise in euro area productivity is associated with an appreciation of the euro by 3% in the long run. The magnitude of this effect suggests that this variable may capture productivity-induced demand effects in addition to the Balassa-Samuelson-type supply-side effects. Likewise, an increase in the real oil price and the euro area's net foreign asset position support the euro vis-à-vis the yen in the long run. The positive effect of the oil price could be expected *ex ante*, given the higher oil dependency of Japan with respect to the euro area. For the government spending variable, the short to medium-term demand effects supporting the currency appear to be more important in the sample period than the more long-term confidence effects, which should rather weigh on the external value of a currency.

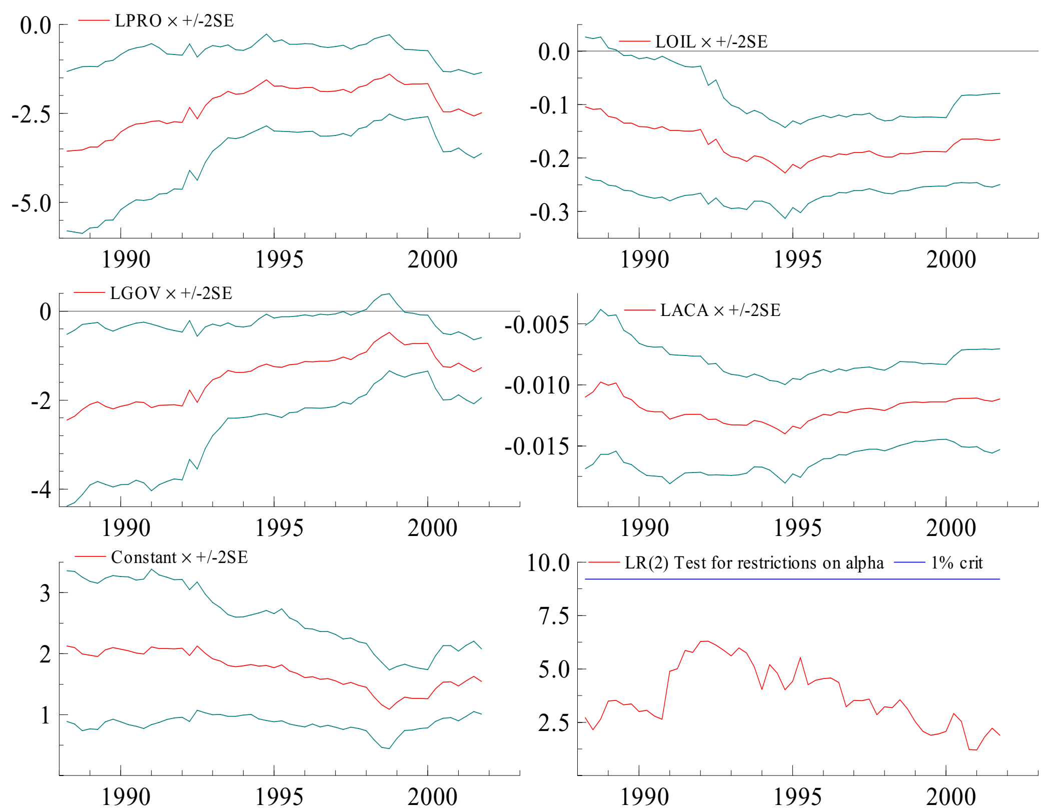
The adjustment term for the real exchange rate is negative and significant, which implies that the real euro-yen exchange rate is indeed one of the variables in the system that adjusts to exogenous shocks. The speed of adjustment is quite fast: following a shock, the half-life of deviations from equilibrium is less than one year.[16] By contrast, the real price of oil and the accumulated current account are individually found to be weakly exogenous with respect to the long-run parameters. The joint hypothesis of weak exogeneity for both the oil price and

[16] The half life is computed as $\log(0.5)/\log(1-\alpha)$, where α is the adjustment term in the equation for the real exchange rate.

the accumulated current account cannot be rejected (tail probability: 0.39). These two variables were thus restricted to be weakly exogenous, and the model was re-estimated under the weak exogeneity restrictions, without having a major impact on the estimates of the coefficients of the long-run vector (specification II).

Stability of the system: Recursive parameter estimates and standard errors for this specification, shown in Chart 5, suggest that the system is reasonably stable.[17] The recursive parameter estimates stay consistently in negative territory for the individual variables and in positive territory for the constant term, as suggested by the cointegration results for the whole sample period. Moreover, for most variables the two standard error bands do not cross the zero line, which implies that the statistical significance of the estimated coefficients is stable over time. A marginal exception constitutes the error band for the relative government spending variable, which shortly crosses the zero-line in the second half of the 1990s, possibly reflecting the theoretical ambiguity of the impact of government spending on the real exchange rate. Finally, the bottom-right chart supports the restrictions on the adjustment terms, reinforcing the proposition that developments in real oil prices and the net foreign asset position are weakly exogenous.

Chart 5. Recursive parameter estimates



[17] The recursive parameter estimates have been carried out by fixing the short-run dynamics at their full-sample values (see Hansen and Johansen 1999).

More information on the characteristics of the model can be derived from the estimated parameters from the moving average representation. These estimates are derived from the maximum likelihood cointegration estimators, as the corresponding parameters are functions of the parameters in the VECM representation. Table 3 displays the orthogonal complements to α, which indicate what variables contribute to each common trend in the system. The table also shows the loadings to these common trends and the long-run impact matrix.

Table 3: The orthogonal complements and the long-run impact matrix

	q	prod	oil	gov	aca
Orthogonal complements to alpha					
$\alpha_{\perp}^{1}$	0.000	0.000	**1.000**	0.000	0.000
$\alpha_{\perp}^{2}$	0.036	**0.999**	0.000	-0.008	0.000
$\alpha_{\perp}^{3}$	0.000	0.000	0.000	0.000	**1.000**
$\alpha_{\perp}^{4}$	0.405	-0.008	0.000	**0.914**	0.000
Loadings to the common trends					
A_1	0.092	-0.016	1.254	0.003	-6.928
A_2	**5.543**	**5.296**	**-10.509**	**-5.936**	**147.53**
A_3	0.037	-0.003	0.067	-0.001	3.103
A_4	1.061	-0.084	1.501	0.423	43.512
C Matrix (long-term impact matrix)					
$\Sigma\varepsilon_{q}$	**0.627**	**0.155**	0.234	-0.040	22.88
	[1.86]	[1.92]	[-0.20]	[-0.43]	[0.99]
$\Sigma\varepsilon_{prod}$	5.531	**5.295**	-10.513	**5.935**	147.094
	[0.77]	[3.11]	[-0.43]	[-3.06]	[0.30]
$\Sigma\varepsilon_{oil}$	0.092	-0.016	**1.254**	0.003	-6.929
	[1.38]	[-1.01]	[5.51]	[0.14]	[-1.52]
$\Sigma\varepsilon_{gov}$	**0.928**	-0.116	1.451	**0.431**	38.663
	[2.47]	[-1.30]	[1.13]	[4.21]	[1.50]
$\Sigma\varepsilon_{aca}$	**0.037**	-0.003	0.067	-0.001	**3.102**
	[2.68]	[-0.92]	[1.44]	[-0.35]	[3.34]

Note: The figures in bold in the first matrix highlight the common trends, and in the second matrix, they emphasise the common trend which has the strongest influence on a variable. In the third matrix, the bold figures indicate statistically significant relationships according to the t-values which are presented in square brackets below.

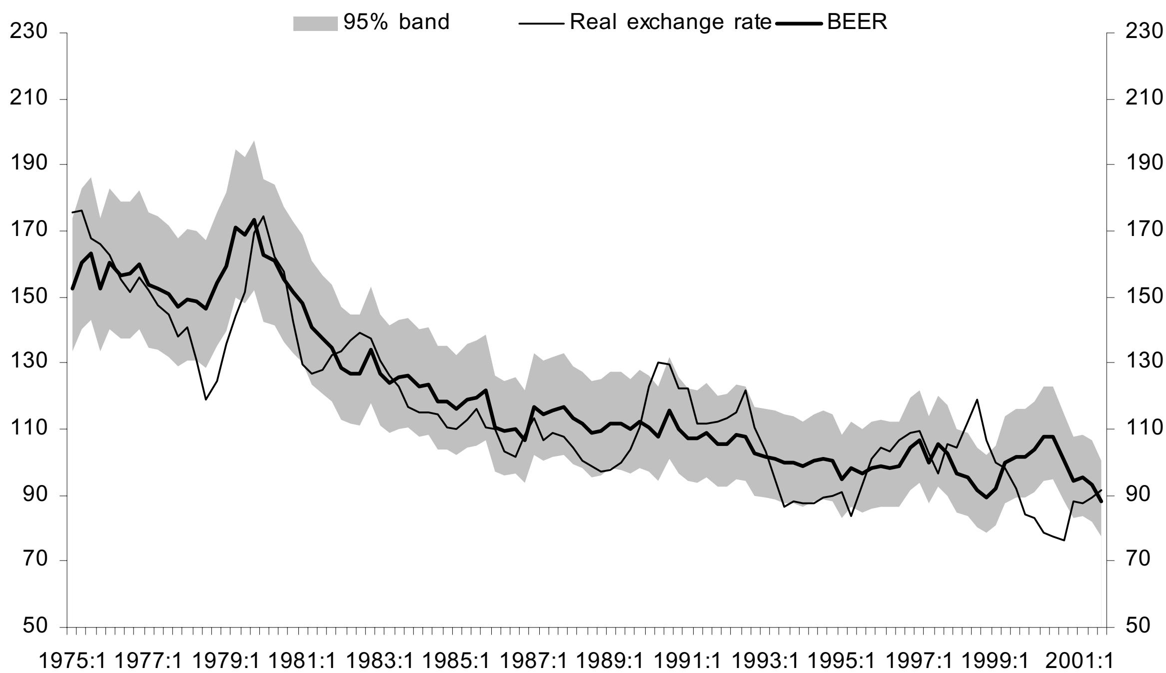
The first rows of Table 3 show the α orthogonal components, which confirm the proposition that the specification explains the real exchange rate, as the common trends are associated with the productivity differential, the real oil price, relative government spending and the accumulated current account, but not with the real exchange rate itself. As illustrated in the second section of the table, showing which variable is most influenced by common trends, the productivity differential is strongly driven by the common trends.

The third matrix – the so-called impact or C matrix – combines this information by multiplying the two matrices mentioned above (see equation (14)) and explains which variables exert a cumulative impact on the real exchange rate. The coefficients of the impact matrix broadly confirm the results of the beta vector in the VECM. The productivity differential, the real price of oil, relative government spending and the relative accumulated current account variable are positively related to the real exchange rate. The productivity variable and the real price of oil are insignificant in this representation, but this does not affect the interpretation of the significance of the coefficients in the cointegration vector, because the estimator of C and its variance also depends on short-run parameters. Shocks on the real exchange rate have a positive and significant impact on productivity, which is a reasonable result since a real appreciation also puts pressure on the domestic economy to improve productivity. Shocks to productivity seem also to have a positive impact on government spending, reflecting possibly a favourable impact on the tax base which gives them more leeway for spending. Finally, the third and the fifth column of the C matrix support the notion that the real price of oil and the accumulated current account are weakly exogenous, since they are only driven by themselves.

3.2.3. Equilibrium exchange rates and error bands

Chart 6 compares the actual real yen-euro exchange rate (Index 1999Q1=100) with the BEER derived from the cointegration vector of specification II. In addition, it shows the two-standard error confidence bands around the estimated BEER. The error bands amount to ±13%, which documents explicitly how difficult it is to determine a precise "equilibrium value" for a currency. Overall, however, the model portrays the real appreciation of the Japanese yen quite well between the mid-1970s and the end-1980s. For the whole sample period, the correlation of the BEER and the actual real exchange rate amounts to more than 85%.

Chart 6: The ("synthetic") yen-euro real exchange rate and its behavioural equilibrium exchange rate (BEER)



In the first half of the 1990s, the yen-euro exchange rate diverged from its equilibrium schedule over protracted periods; in the medium term, however, it always reverted towards this "centre of gravity" – the equilibrium schedule. In more detail, the sharp depreciation of the yen in 1989/90 is only partially captured by the model. As a result, the yen-euro real exchange rate touched the lower bound of the confidence band, indicating that the yen was against the euro below its equilibrium value ("undervalued") up to 1993. The BEER estimate also implies that the sudden appreciation of the yen in 1993 may have reflected initially a reversion of the yen to "equilibrium". Between the second half of 1993 and the first half of 1995, however, the yen was according to these computations by almost the same amount above its equilibrium. Between 1996 and the first half of 1998, the yen fluctuated closely around its equilibrium.

Subsequently, the pattern of the equilibrium schedule of the real yen-euro exchange rate and its actual counterpart becomes rather peculiar. The two series indeed appear to be negatively correlated over the last four years under consideration, but the gravity forces towards the equilibrium schedule still seem to be in place. In 1998, for instance, the yen experienced a strong depreciation vis-à-vis the euro legacy currencies, which catapulted the (real) yen-euro exchange rate shortly outside the confidence bands, but reverted forcefully to its "equilibrium" in 1998Q4 and 1999Q1. Following the launch of the euro, the euro depreciated between the first quarter of 1999 and the fourth quarter of 2000 by roughly 27% vis-à-vis the Japanese currency (in real terms). Over the same period, however, Japan was lagging behind in terms of productivity developments and the surge in oil prices, which traditionally weighed on the yen, failed to support the euro.

These effects may have been partly offset by remarkable current account surpluses in Japan in this period, while the euro area recorded a modest deficit in its current account balance. On the other hand, the net foreign asset positions of both the euro area and Japan were in positive territory according to their accumulated current account balances. As a result, the decline of the euro vis-à-vis most major currencies should have increased the value of foreign assets relative to domestic output. Overall, the BEER rose in this period, implying an equilibrium appreciation of the euro vis-à-vis the Japanese yen, while the Japanese currency appreciated to its highest level in the last 25 years (in the fourth quarter of 2000). Indeed, according to these estimates, the real euro-yen exchange rate was in 2000 persistently outside the confidence band. Since the end of 2000, the euro has appreciated significantly against the yen, reversing to a large extent its earlier depreciation. At the same time, the depreciation of the yen vis-à-vis the US dollar had a positive effect on the Japanese net foreign asset position, which strengthened the BEER of the yen vis-à-vis the euro, and led to a narrowing and finally to a correspondence of the actual exchange rate and the "equilibrium schedule" towards the end of 2001.

Finally, regarding the general appreciation trend of the Japanese yen over the entire sample period it is instructive to use the estimation results to assess the importance of the various fundamental variables in bringing about this strengthening. From 1975Q1 until 2001Q4 the Japanese yen appreciated by 65% vis-à-vis the ("synthetic") euro in real terms (log percentage changes). While oil price and relative government spending can explain only little of this trend, the relative productivity variable and the relative net foreign asset position both moved in favour of Japan, thus strongly supporting the Japanese yen. The estimated coefficients from the long-run cointegrating relationship suggest that those two factors account for roughly 90% of the yen's (equilibrium) appreciation. Most of the appreciation over the sample period was actually concentrated in the 1980s, with the Japanese currency appreciating in real terms by more than 30% between 1980Q1 and 1988Q4 alone. During that period, however, the influence of the oil price variable was rather strong. Thus, the rise of the yen appears to be driven by a number of different factors, with the traditional productivity explanation – often cited as the main explanation[18] – being an important but by no means the only factor.

4. Conclusions

This paper provides an extensive empirical analysis of the "synthetic" euro-yen exchange rate over the period 1975–2001. Using cointegration analysis, we find a consistent and significant relationship between the real exchange rate and relative productivity, net foreign assets, relative government spending and the terms of trade. A fairly rapid mean reversion of the exchange rate to its equilibrium is also detected over the sample period. Overall, the "equilibrium" euro-yen exchange rate derived from this model tracks the development of the actual exchange rate rather well between the mid-1970s and the end of the 1990s. The estimates suggest that the substantial appreciation of the yen over this period – especially pronounced during the 1980s – was not only driven by productivity differential, but also the accumulation of foreign assets and developments in oil prices. Towards the end of the period under consideration, the relationship between the equilibrium schedule of the real yen-euro exchange rate and its actual counterpart has become rather peculiar. The two series indeed appear to be negatively correlated in this episode, but the gravity forces towards the equilibrium schedule still seem to be in place. In particular, the model suggests, that the euro appreciation vis-à-vis the yen in 2001 has represented largely an equilibrium correction of its depreciation in 1999/2000.

18 See, for instance, Yoshikawa (1990).

Annex 1: "Error bands" for the BEER

Following Johansen (2001) we express the common non-stationary trends as linear combinations of the observable data instead of the unobservable shocks. Disregarding for simplicity the deterministic part, this is done by multiplying (12) by $\alpha'_\perp$ and cumulating:

(A.1.) $$\alpha'_\perp \Delta y_t = \sum_{i=1}^{k-1} \alpha'_\perp \Gamma_i \Delta y_{t-i} + \alpha'_\perp \sum_{i=1}^{t} \varepsilon_i \, .$$

Using the definition of $\Gamma = I - \Sigma_{i=1}^{k-1} \Gamma_i$, one can solve for the common trends as a function of levels, differences and initial values:

(A.2.) $$\begin{aligned} \alpha'_\perp \sum_{i=1}^{t} \varepsilon_i &= \alpha'_\perp \Gamma y_t - \sum_{i=1}^{k-1} \alpha'_\perp \Gamma_i (y_{t-i} - y_t - y_{-i}) - \alpha'_\perp y_0 \\ &= \alpha'_\perp \Gamma y_t + Z_t + A \end{aligned}$$

where Z_t is an I(0) process and A is a function of initial values. Looking at the definition of

(A.3.) $$C = \beta_\perp (\alpha'_\perp \Gamma \beta_\perp)^{-1} \alpha'_\perp$$ we get:

(A.4.) $$C \sum_{i=1}^{t} \varepsilon_i = C\Gamma y_t + \beta_\perp (\alpha'_\perp \Gamma \beta_\perp)^{-1} (Z_t + A) \, .$$

Then a simple non-stationary – stationary decomposition is given by:

(A.5.) $$y_t = C\Gamma y_t + (I - C\Gamma) y_t \, .$$

One can then use the variance of the stationary part to derive bands around the nonstationary component. Taking a unit vector to select the first variable in the system: $y_{1t} = u'_1 y_t$, then

(A.6) $$\begin{aligned} u'_1 y_t &= u'_1 C\Gamma y_t + u'_1 (I - C\Gamma) y_t \qquad \text{or:} \\ y_{1t} - u'_1 C\Gamma y_t &= u'_1 (I - C\Gamma) \bar{\beta} \beta' y_t \end{aligned}$$

Where $\bar{\beta} = \beta(\beta\beta)^{-1}$, such that $\bar{\beta}\beta' = I$. Then one can plot the following bands:

(A.7) $$u'_1 C\Gamma y_t = \pm 2 \sqrt{\left((I - C\Gamma)\bar{\beta} \, \mathrm{Var}(\beta' y_t) \, \bar{\beta}'(I - C\Gamma)'\right)_{11}} \, .$$

Annex 2: The determinants of the JPY/DEM exchange rate

In order to conduct a robustness check, we estimated the model discussed above also for the JPY/DEM currency pair. As was the case for the JPY/EUR, the real interest rate differential did not play a role. Both relative productivity and the government expenditure differential have a very similar magnitude as for the JPY/EUR, while neither the real price of oil nor the net foreign assets turned out to be significant. The main results are summarised in Table A, both for the whole sample and for a restricted sample up to 1998q4, when the D-mark was substituted by the euro. For estimation in the later sample we used the JPY/DEM rate implied by the corresponding euro rate and the irrevocably fixed DEM/EUR parity.

The model was estimated with four lags, including a dummy variable for government expenditure to deal with two outliers of opposite sign, in 1986q1 and 1990q3. Since this dummy sums to one, it does not affect the asymptotics of the trace test. The standard specification tests indicate that the residuals are white noise, the variables are significant in the long-term relationship and their signs are in line with the theoretical reasoning described in section 2. Furthermore, as stated above, the coefficients are very close in magnitude to the coefficients of productivity and government expenditure for the JPY/EUR currency pair, as is evidenced by comparing Table 2 in the text with Table A in this annex.

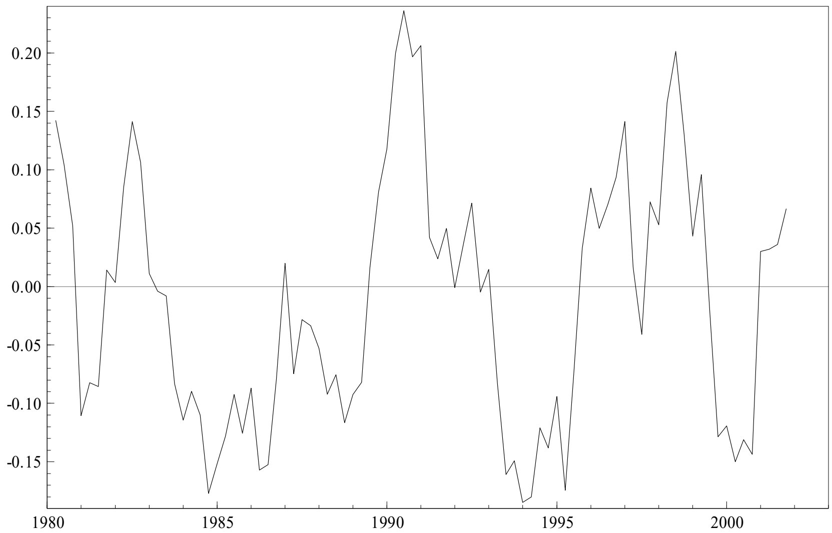
Table A: VECM results

	1980q2-1998q4		1980q2-2001q4	
	Long-term coefficients *(t statistic)*	Adjustment terms *(t statistic)*	Long-term coefficients *(t statistic)*	Adjustment terms *(t statistic)*
q	1.000	-0.120	1.000	-0.190
	-	*(-2.300)*	-	*(-3.438)*
prod	-2.815	0.044	-2.152	0.043
	(-2.923)	*(2.916)*	*(-2.8119)*	*(2.656)*
gov	-1. 112	0	-0.921	0
	(-2.715)	-	*(-4.632)*	-
constant	10.094		9.602	
	(2.806)		*(2.725)*	
Summary Statistics:				
Test on restrictions *(p-value)*	1.603 *(0.20)*		1.4760 *(0.22)*	
LM(1-5) test	F(45,167) = 0.87 *(0.70)*			
Cointegration test*				
		Standard critical values		
		5%	1%	
	37.72	34.91	84.45	
	18.19	19.96	24.60	
	3.92	9.24	12.97	

* Test for the longer sample period. Constant restricted to the cointegrating space. Standard critical values for the model with a constant restricted to the cointegrating space as suggested by Osterwald-Lenum (1992).

Choice of cointegration rank: The trace tests indicate the presence of one cointegration relationship (based on standard critical values). The hypothesis of one cointegration relationship is confirmed by looking at the magnitude of the roots of the system, which displays an eigenvalue equal to 1.018 and a smaller one at 0.916 when unrestricted. Once the rank restriction and the weak exogeneity of government expenditure differentials are imposed, the third largest root drops to 0.759, supporting the rank choice. The graph of the cointegration vector, which depicts $\beta' y_t$, is displayed in Chart A. The cointegration vector appears indeed to be stationary, and this additional evidence supports the finding of a rank equal to one suggested by the standard trace test.

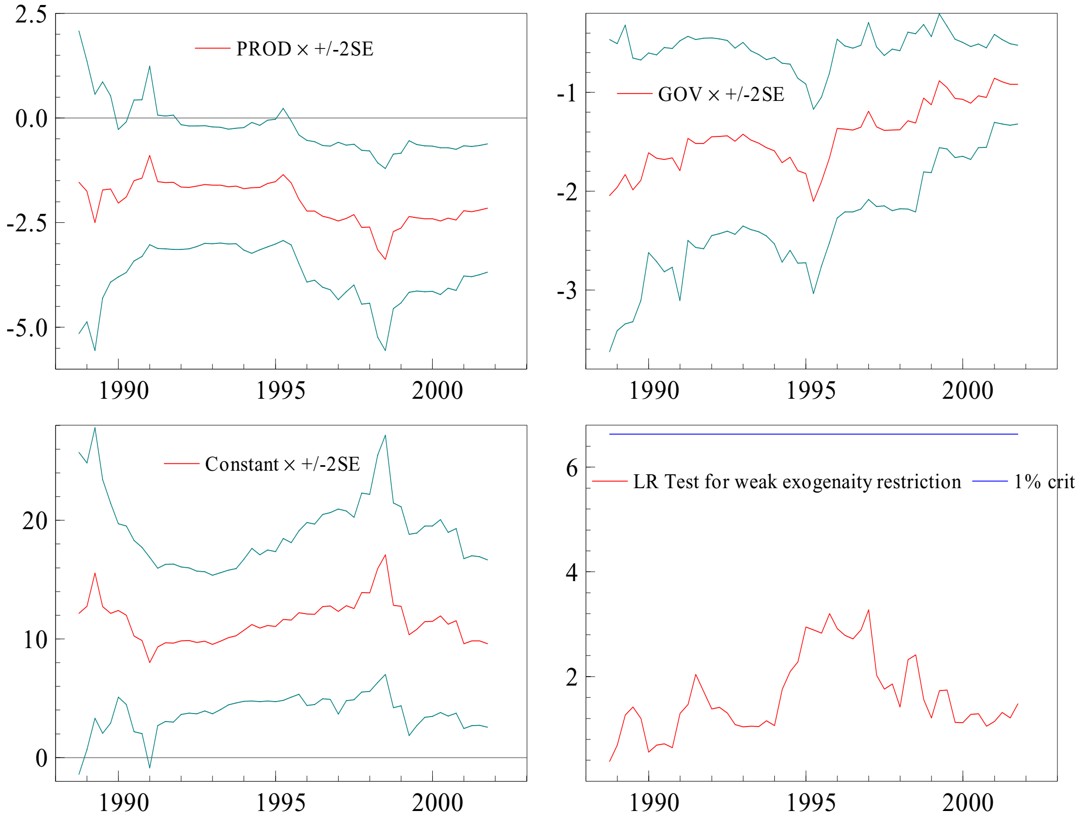
Chart A: Cointegration vector ($\beta' y_t$)



The adjustment term for the real exchange rate is negative and significant, which implies that the real JPY/DEM exchange rate is indeed one of the variables in the system that adjusts to exogenous shocks. The speed of adjustment is rather fast: following a shock, the half-life of deviations from equilibrium is about one year. In contrast to what is found for the JPY/EUR, the government expenditure differential is weakly exogenous with respect to the long-run parameters.

Recursive parameter estimates and standard errors for this specification, shown in Chart B, suggest that the system is reasonably stable. The recursive parameter estimates stay consistently in negative territory for both individual variables and in positive territory for the constant term, as suggested by the cointegration results for the whole sample period. Moreover, for most variables the two standard error bands do not cross the zero line, which implies that the statistical significance of the estimated coefficients is stable over time.

Chart B. Recursive parameter estimates



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