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Abstract:

This paper presents an empirical analysis of the medium-term determinants of the euro effective exchange rate. The empirical analysis builds on synthetic quarterly data from 1975 to 1998, and derives a Behavioural Equilibrium Exchange Rate (BEER) and a Permanent Equilibrium Exchange Rate (PEER). Four different model specifications are retained, due to the difficulties encountered in specifying an encompassing model. Results indicate that differentials in real interest rates and productivity, and (in some specifications) the relative fiscal stance and the real price of oil, have a significant influence on the euro effective exchange rate. Assessing the existence and the extent of the over- or undervaluation of the exchange rate is not straightforward, since these different specifications often lead to contrasting findings. However, all four models point unambiguously to the undervaluation of the euro in 2000, although the extent of this undervaluation largely depends on the specification chosen.

JEL classification system: F31, F32.

Keywords: *euro, equilibrium exchange rates, cointegration analysis, Gonzalo-Granger decomposition, fundamentals, BEER, PEER.*

Non-technical summary

The exchange rate of the euro has experienced a marked decline since its launch in 1999. In effective terms, the depreciation of the euro amounted to roughly 17½% between the first quarter of 1999 and the fourth quarter of 2000. This evolution of its external value has raised concerns that the exchange rate of the euro might have moved out of line with fundamentals.

Such an assertion requires a measure of the "equilibrium" exchange rate as a benchmark against which the actual development of the exchange rate can be gauged. Although it is widely accepted that providing a precise estimate of the equilibrium level of exchange rates and thus of the over- or undervaluation of a currency is far from straightforward, a number of empirical models based on economic fundamentals have shown that they can track the evolution of the actual exchange rate rather well. When the discrepancies between the estimated equilibrium values and the actual exchange rate become extraordinarily large, such models may serve to suggest the direction of misalignments, and thus support the qualitative judgement that exchange rates are out of line with economic fundamentals. In other words, while we may find clear indications that the exchange rate is inconsistent with fundamentals, we are not in a position to declare a certain value to be the correct equilibrium value.

This paper presents an empirical analysis of the medium term determinants of the effective exchange rate of the euro, using only one particular approach: the so-called Behavioural Equilibrium Exchange Rate (BEER), which effectively involves reduced form modelling of the exchange rate based on standard cointegration techniques. Quarterly data spanning the period 1975 to 1998 and subsequently extended to 2000 have been employed in the empirical analysis. In view of the short history of EMU, "synthetic" data was compiled for the euro area by aggregating the data of the individual euro area countries using trade weights. Based on standard economic theory, the paper tests whether the real interest rate differential, the productivity differential, the relative fiscal stance, time preferences, the real price of oil and the (accumulated) current account have a significant influence on the effective exchange rate of the euro. The econometric analysis starts by analysing the stochastic properties of the data, and finds that most of the variables are non-stationary. Accordingly, it proceeds by estimating vector-error correction models. While it was not possible to derive a unique model, encompassing all other specifications, four well-specified models including subsets of the variables have been identified, which yield unique cointegration vectors. In all of these models, the variables have the theoretically expected signs and are statistically significant.

The analysis suggests that the euro effective exchange rate is mainly driven by the productivity differential, either measured directly, as the ratio of real GDP to the total number of employees, or indirectly, as the relative price differential between traded and non-traded goods. The more direct measure proxies trends in labour productivity in the economy while the latter, more indirect variable, tries to capture the different productivity trends in traded and non-traded goods sectors separately, and may therefore embody Balassa-Samuelson effects. Two of the models also include the real price of oil, while the other two incorporate the real expected long-term interest rate differential. In the two models where the productivity differential was measured indirectly by the relative price variable, the government spending variable was also significant. In contrast, the inclusion of the net foreign asset position did not lead to a consistent model, and both the proxy for the rate of time preference and the short-term interest rate differential proved to be insignificant. Overall, therefore, the euro appears to be mainly affected by productivity developments, (expected) real interest rate differentials, and terms of trade shocks due to the oil dependence of the euro area.

The analysis is complemented by decomposing the estimated BEER into a transitory and a permanent component, following the procedure proposed by J. Gonzalo and C. W. Granger, which yields the so-called Permanent Equilibrium Exchange Rates (PEER). The PEERs, which are smoother because they are purged of transitory effects, turned out to be quite similar to the BEERs obtained from the cointegrating vectors. Both the BEERs and the PEERs indicate that the euro was close to its fundamental value (or slightly overvalued) in the seventies and in the first half of the nineties. The real effective exchange rate of the euro experienced a depreciation of 17% in nominal effective terms from the first quarter of 1999 to the last quarter of 2000. The estimated models attribute some of that depreciation to a decline in the equilibrium rate, but they also unanimously indicate an undervaluation by the end of 2000. The extent of this undervaluation, however, largely depends on the model specification and is surrounded by a significant margin of error.

However, periods when the majority of the models agree on a notable misalignment in the same direction are rather infrequent. The observation that a sizeable amount of undervaluation was signalled unanimously by all four estimated models in 2000 provides support to the conclusion that the euro was out of line with fundamentals at that time.

"I know that the exchange rate of the euro […] does not yet reflect the fundamentals […], but I am not in a position to declare a certain value to be the right value."
Wim Duisenberg, press conference, Frankfurt, 8 June 2000.

I. Introduction

The exchange rate of the euro has been characterised by a marked decline since its launch in 1999. In October 2000 the euro was quoted against major currencies at its lowest levels recorded thus far, which implied a depreciation of more than 30% against both the US dollar and the Japanese yen compared with the beginning of 1999. In effective terms, i.e. weighted against the currencies of the thirteen most important trading partners of the euro area, the decline of the euro amounted over the same period to roughly 23%. At that time, these movements in the exchange rate raised concerns about the risks they may pose for the world economy.

Any assertion on disequilibrium in the foreign exchange market requires a measure of the "equilibrium" exchange rate as a benchmark against which the actual development of the exchange rate is gauged. Although it is widely accepted that providing a precise estimate of the equilibrium level of exchange rates and thus of the over- or undervaluation of a currency is far from straightforward, a number of empirical models based on economic fundamentals have shown that they can track the evolution of the actual exchange rate rather well.[1]

This paper examines the role played by economic fundamentals in explaining the behaviour of the euro real effective exchange rate. The estimated models have a medium-term focus and as the application of appropriate econometric methods requires a relatively large number of observations, quarterly data spanning a long period were used. Hence, the construction of "synthetic" historical time series has become necessary in view of the fairly short period for which data on the exchange rate of the euro and its fundamentals are available. While many empirical approaches to modelling equilibrium exchange rates have been documented in the literature,[2] the analysis in this paper focuses on the computation of Behavioural Equilibrium Exchange Rates (BEER) based on a reduced form specification, which links the effective exchange rate of the euro to a broad set of economic fundamentals. Their interaction is analysed empirically by applying standard cointegration techniques and by decomposing the cointegrated time series into their permanent and transitory components.[3] The results suggest

1 Following the influential paper by Meese and Rogoff (1983), the acid test for such models was to beat the naive random walk in terms of forecasting properties. Clostermann and Schnatz (2000) as well as MacDonald and Marsh (1997), for instance, show that with models based on a similar methodology as ours it is possible to outperform the naive random walk model. While the evaluation of forecast performance is beyond the scope of the present paper, it serves to indicate a convenient starting point.

2 For overviews, see, for instance, Driver and Westaway (2001) and Mac Donald (2000).

3 This is the so-called Permanent Equilibrium Exchange Rate (PEER); see, for instance, Clark and MacDonald (2000) for a description of this concept.

that the real long-term interest rate differential, productivity-related variables and, to a lesser extent, fiscal variables and oil prices can be identified as fundamental determinants of the euro exchange rate. All of these models lead to the conclusion that the euro was below its equilibrium level at the end of 2000.

A word of caution is necessary when one considers the evidence provided by the models estimated in this paper, which applies to all fundamental-based models of the real exchange rate: the degree of uncertainty surrounding the estimation results is rather large. However, these models may serve to suggest the direction of misalignments and support the qualitative judgement that exchange rates are out of line with economic fundamentals when the discrepancies between the estimated equilibrium values and the actual exchange rate become extraordinarily large. In other words, while there is evidence that the exchange rate is out of line with fundamentals, this is not enough to declare a certain value to be the correct equilibrium value.

The paper is organised as follows: section II presents an overview of fundamental-based models of real exchange rate determination and of some related empirical literature, placing special emphasis on studies that include the euro in a BEER/PEER framework. Section III describes the data and explains the choice of variables. The core of the analysis is in section IV, where the econometric methodology is discussed and the empirical results are presented. The final section summarises the main conclusions.

II. Fundamental-based models of the real exchange rate: an overview

a. Approaches to modelling the medium-term determinants of the exchange rate

For many empirical studies on exchange rates, the starting point has been the purchasing power parity (PPP) doctrine, which claims that the exchange rate is determined by the relative developments of domestic and foreign prices, thereby suggesting that the equilibrium real exchange rate is a constant. However, it is well documented in the literature that the real exchange rate is either found to be non-stationary or, when found to be mean-reverting in studies using a very long sample span or applying panel data analysis, its adjustment speed to the equilibrium path mapped out by relative prices is very slow, so that prolonged deviations from its equilibrium cannot be explained on the basis of this concept.

Owing to these limitations, a majority of recent studies use more sophisticated approaches. In particular, they explicitly model the equilibrium exchange rate as a function of real economic fundamentals, thereby allowing for a time-varying equilibrium path of the real exchange rate. The underlying theoretical framework is usually broadly consistent with the tradition of the macroeconomic balance approach. A prominent methodology in this context is the so-called 'fundamental equilibrium exchange rate' (FEER) model, advocated by Williamson (1994).

The FEER is defined as the exchange rate consistent with internal and external balance. Internal balance is obtained when a country is operating at a level of output consistent with full employment and low inflation. External balance can be characterised by a sustainable current account position as reflected by the underlying and desired net capital flows, which depend on net savings that are, in turn, determined by factors such as consumption smoothing and demographic factors. The FEER approach can be characterised as normative in the sense that it delivers an equilibrium exchange rate consistent with 'ideal' economic conditions.[4] The so-called NATREX models are from a theoretical point of view related to this concept (Stein 1994, 1999). In the end, the NATREX is usually also derived from reduced-form equations which are, however, based on dynamic stock-flow models.[5] A feature of this approach is that net foreign assets and the capital stock are assumed to be at their steady-state level.

This paper follows a related strand of literature by focusing on the so-called BEER ('behavioural equilibrium exchange rates') and PEER ('permanent equilibrium exchange rates') approaches proposed by Clark and MacDonald (1999, 2000). According to Driver and Westaway (2001), BEERs can be categorised as "current and cyclical equilibrium exchange rates", since their computation is based on the current levels of the fundamental factors. Moreover, they may also include (cyclical) variables that can have a persistent effect on the exchange rate, which should, however, wash out over time. The transition from the current or cyclical to the medium-term perspective has been accomplished by decomposing, with further statistical refinements, the equilibrium exchange rate derived under the BEER methodology into permanent and transitory components, thereby deriving the PEER. This methodology also allows for a time-varying equilibrium of the real exchange rate, but places less normative structure on the model and on the computations.[6] In these models, movements in the real exchange rate are assumed to be mainly determined by relative sectoral productivity differentials and the outstanding stock of net foreign assets. Since the theoretical background to this model has been formally explained in detail in many papers, it will not be elaborated further.[7]

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4 See Driver and Wren-Lewis (1999) for a comprehensive application of the FEER methodology.

5 A notable exception is Detken et al. (2001), who compute equilibrium exchange rates of the euro based on the NATREX approach estimating a structural model.

6 Therefore, the equilibrium concept in this paper is rather statistical than theoretical, as it is not required for the explanatory variables themselves to be at their equilibrium or steady-state levels; See Detken et al. (2001) for these remarks.

7 The reader may refer to MacDonald (1997), Clostermann and Friedmann (1999) and Clostermann and Schnatz (2000), for example.

b. Recent empirical studies applying the BEER/PEER approach to the euro

Given the existence of several comprehensive surveys on empirical exchange rate economics,[8] this section only covers a selection of recent papers that use the BEER/PEER framework to study the euro exchange rate in bilateral or effective terms.

One of the most ambitious attempts to address the issue of equilibrium exchange rate determination has been presented by Alberola et al. (1999). They rely on cointegration techniques for individual currencies as well as for a panel of currencies and infer the long-run properties of various real effective exchange rate indices. However, their model depends on a very limited set of fundamentals: net foreign assets and relative sectoral prices as a proxy for Balassa-Samuleson-related productivity differences. A critical feature of the study by Alberola et al. is that the panel of currencies used involves both the "synthetic" euro effective exchange rate and the effective exchange rates of the major EMU legacy currencies, which violates the independence assumption that underlies panel estimation and inference. Moreover, although the reported parameters have the expected signs, it is difficult to infer how important the individual fundamentals are and whether they are indeed determinants of the real exchange rate trends, as not enough details are given in the test statistics. The paper of Roeger and Hansen (2000) shares many of the above-mentioned drawbacks. No information on the statistical significance of the variables claimed to explain long-term movements in the exchange rate is reported, and the results of the econometric computations are sometimes replaced by rather pragmatic assumptions if they do not provide the desired results.

Lorenzen and Thygesen (2000) have elaborated a very complete empirical assessment. They focus on the bilateral euro exchange rate against the US dollar and compute a long-run, a medium-run and a cyclical equilibrium exchange rate. For the long-run equilibrium exchange rate they go beyond many other studies by including, in addition to regularly employed variables like the net foreign asset position or relative sectoral prices, the following determinants: the dependency ratio (as a proxy for the propensity to save and consume), expenditure on R&D (as an additional proxy for the Balassa-Samuelson effect) and a risk premium (which turns out to be insignificant in the empirical analysis). However, this has the drawback that some of these data are only available on an annual basis, so that in order to obtain enough observations the sample has had to be extended back to 1960, thus encompassing various exchange rate regimes. While such a procedure could introduce a bias into the results, it is a common practice.[9]

8 See for example Froot and Rogoff (1994), Frankel and Rose (1995), MacDonald (1995, 2000), and Stein (2001).

9 For instance, Faraquee (1995) studies the effective exchange rate of the yen and dollar for the 1950-90 period and finds that the yen is cointegrated with productivity proxies (measured either as relative prices or as productivity in the tradable sector), while the US dollar is cointegrated with relative prices and the net foreign asset position.

Clostermann and Schnatz (2000) focus on the fundamental determinants of the bilateral euro-dollar exchange rate. Their analysis basically covers the period of floating exchange rates and includes an internal price ratio differential (as an indirect productivity proxy) and the real interest rate differential, the real oil price and the relative fiscal position to explain the exchange rate of the euro against the dollar. The authors show that there is a stable long-run relationship between the real exchange rate and these fundamentals. The paper by Makrydakis et al. (2000) studies an effective exchange rate of the euro against ten countries between the first quarter of 1980 and the second quarter of 1999. They find cointegration between the real effective exchange rate, a productivity variable (measured as real GDP per employee)[10] and the real interest rate differential, while the net foreign asset position proves to be insignificant. Moreover, the fitted values are decomposed into their permanent and transitory components, which enables the derivation of a PEER.

The empirical analysis undertaken in this paper builds on the last two studies and extends them. First of all, it applies a consistent data set for the real effective exchange rate of the euro against twelve countries[11] over the sample period of 1975 – 2000. Secondly, the relevance of all variables suggested in both papers is evaluated in the present paper.

III. Construction of the data, motivation and definition of the variables

a. Construction of synthetic time series

In view of the short history of the European Monetary Union, a medium-term analysis of real effective exchange rate developments requires the construction of historical time series for the euro area by aggregating the data of the individual euro area participating countries. These 'synthetic' time series are computed on the basis of quarterly data covering the 1975:1-2000:4 period, which broadly corresponds to the period of floating exchange rates following the collapse of the Bretton Woods system, allowing for a short phase of adjustment to the new circumstances in the foreign exchange markets.

Each time series for the euro area (X^E) has been computed as a geometric weighted average of the individual euro area countries series, using the weights (w_j) of each euro area participating country j in total manufacturing trade of the euro area:[12]

$$X_t^E = \prod_{j=1}^{11} (X_t^j)^{w_j}$$

[10] This productivity variable was proposed in Clostermann and Friedmann (1998).

[11] Those included in the narrow group in the European Central Bank's computations of effective exchange rates, i.e. Denmark, Sweden, United Kingdom, Australia, Canada, Japan, Norway, Switzerland, United States, South Korea, Hong Kong and Singapore.

[12] The euro area includes data for Greece, treating Greece as a member of the euro area over the entire period, while the weights for Luxembourg and Belgium are merged. See the appendix A1 for the weighting scheme.

The partner countries of the euro area are its twelve major trading partners. They encompass the EU countries outside the EMU (United Kingdom, Sweden and Denmark), the United States, Japan, Canada, Australia, Switzerland, Norway, Hong Kong, Singapore and Korea. The weights (g_i) for compiling these data are based on manufacturing trade flows data averaged over the 1995-1997 period and accounting for third market effects:[13]

$$X_t^P = \prod_{i=1}^{12} (X_t^i)^{g_i}$$

Correspondingly, the 'synthetic' nominal effective exchange rate (E^E) of the euro is given as

$$E_t^E = \prod_{j=1}^{11} \prod_{i=1}^{12} ((E_t^{ij})^{w_j})^{g_i}$$

where E^{ij} is the exchange rate for the partner currencies i against each euro legacy currency j (e.g. US dollar per D-Mark), which implies that an increase in E^E reflects an appreciation of the (synthetic) euro in effective terms.

The real effective exchange rate is defined as the nominal effective exchange rate adjusted for movements in consumer price indices at home and abroad:

$$Q_t^E = E_t^E \frac{P_t^E}{P_t^P}$$

b. Motivation and definition of the variables

Drawing on the empirical analyses using the BEER/PEER approach, the following set of economic fundamentals were analysed in the present study:

- ***Productivity differentials***

The impact of the productivity differential on the real exchange rate is expected to follow the well-known Balassa-Samuelson doctrine, which states that relatively larger increases in productivity in the traded goods sector are associated with a real appreciation of the currency of a country.

Two alternative expressions – one indirect and another more direct – were used to take diverging productivity trends into account: the relative price differential between traded and non-traded goods at home and abroad (INT), and the total labour productivity differential (PRO). The first is a indirect proxy which is widely used to capture the effect of productivity increases in the traded goods sector.[14] It is computed as the relative price (consumer to

[13] The import weights are simple shares of each partner country in total euro area imports from the partner countries. Exports are double weighted in order to account for third market effects, so as to capture the competition faced by euro area exporters in foreign markets from both domestic producers and exporters from third countries. For the methodology see ECB (2000).

[14] See, for example, Chinn (1999), Clark and MacDonald (1999), Clostermann and Schnatz (2000), Kakkar and Ogaki (1999), MacDonald (1999).

wholesale price indices) differential between the euro area and its most important trading partners.

$$\mathrm{INT}_t^E = \left(\frac{P_t^{E,NT}}{P_t^{E,T}} / \frac{P_t^{P,NT}}{P_t^{P,T}}\right)$$

The second definition of relative productivity, measured as the total labour productivity differential between the euro area and abroad, was employed as a more direct alternative to the INT variable. Productivity was measured as real GDP (Y) divided by the number of employees (EM):

$$\mathrm{PRO}_t^E = \left(\frac{Y_t^E}{EM_t^E} / \frac{Y_t^P}{EM_t^P}\right)$$

Sectoral productivity differentials focusing on the manufacturing sector would be a better direct proxy, but on a quarterly basis there are no reliable data available to compute them.

Both productivity variables have also some drawbacks, however: firstly, PRO and INT are not necessarily equivalent. For instance, an increase in the productivity in the sector of traded goods of the euro area triggers a rise in both variables; in contrast, a rise in the productivity in the non-traded goods sector leads, ceteris paribus, to a fall in INT but to an increase in PRO. Only if the income elasticity of non-traded goods was high enough, the additional domestic wealth generated by the productivity increase would offset the relative price adjustment. Nevertheless, in principle it could be expected that both variables evolve in a similar way, specifically when catching-up processes are taking place.[15] Therefore, it is assumed that both variables have a positive impact on the real exchange rate. Secondly, changes in taxes (in particular value added taxes) which have an impact on consumer prices may obscure the information implied in this variable on productivity trends. Thirdly, the analyses of more disaggregated price data indicate that the effect of the price changes in the non-traded good sector seems to be rather small (Duval 2001). In principle, therefore, it would be better to construct price indices from industries producing clearly traded or non-traded goods. For the sample under consideration, however, the lack of data availability precluded this.

- ***Net foreign assets***

The inclusion of the outstanding stock of foreign assets as a determinant of the real exchange rate follows portfolio-balance considerations. For instance, a deficit in the current account creates an increase in the net foreign debt of a country, which has to be financed by internationally diversifying investors. However, for the associated adjustment of their portfolio structure, they demand a higher yield. At given interest rates, this can only be accomplished through a depreciation of the currency of the debtor country. In addition, the

[15] Chinn (1999), for instance, finds relative prices and income per capita to be highly correlated in a panel cointegrating study for OECD countries.

balance of payments channel assumes that a current account deficit accumulates net foreign debts for which interest has to be paid. The interest payments need to be financed by an improvement of the trade balance. This, in turn, requires a depreciation of the currency, strengthening the international price competitiveness of the country and increasing the attractiveness of its exports, the proceeds of which are used to service the higher interest payments.[16] The variable ACA, which measures the accumulated current account position as a percentage of GDP, was included in the list of variables. This is a rather distorted proxy for net foreign assets, because it ignores the effects of debt reduction and forgiveness and valuation issues. Furthermore, in the case of the euro area, the constructed cumulated current account measure does not correct for intra-euro area current account positions. Nonetheless, this is the best measure available for the longer-term horizon analysed in this paper. In fact, it turns out that this variable cannot be employed successfully in the estimated model. This result could be explained by the distortion introduced by the measurement problem, and should be interpreted with caution in evaluating the impact of net foreign assets on the exchange rate.

- ***Fiscal position***

In this context, the relationship with the fiscal balance is also interesting, as it constitutes one of the key components of national savings. According to Frenkel and Mussa (1988), a fiscal tightening causes a permanent increase in the net foreign assets position of a country and, consequently, an appreciation of its equilibrium exchange rate in the longer term, provided that the fiscal consolidation is considered to have a permanent character. Accordingly, the relative expenditure ratio between the euro area and abroad (GOV) has also been included. However, a positive impact of government spending on the real exchange rate in the short term could stem from demand-side effects. First, if the private sector lowers its demand for goods less than the increase in government spending, the positive demand shock could affect the real exchange rate via relative prices. Secondly, if the marginal propensity of the public sector to spend on non-traded goods is high (which is plausible), the price of non-traded goods would rise, thus also leading to an appreciation of the real exchange rate.[17] In the longer term, however, higher government spending most likely undermines confidence in a currency, because it could be followed (or accompanied) by distorting taxes and thus have on balance a negative impact on economic growth and the real exchange rate. Furthermore, the

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[16] See Fischer (2000), however, who shows in a NATREX framework that changes in the net foreign asset position in reaction to exogenous shocks (like changes in preferences) could be offset by adjustments in the capital stock. Since changes in the net foreign asset position and changes in the capital stock have countervailing effects, the long-run impact on the real exchange rate may be ambiguous. See also Detken et al. (2001), who argue that an accumulation of net foreign reserves can be associated with a depreciation of the domestic currency in the medium run, but trigger an appreciation in the long run.

[17] See Dibooğlu (1996) for this argument, who also finds some evidence for such a relationship for Germany, Japan and Italy.

impact of the gov variable could be ambiguous, due to the effect of higher government spending on real interest rates. However, this effect is controlled for in the empirical application in this paper by including the real interest rate differential in the empirical equation. As an alternative to government consumption, the total (public and private) consumption to GDP differential in the euro area and abroad (CON) has been incorporated in the model in order to analyse whether time preference effects have an impact on the real exchange rate.

- ***Terms of trade shocks***

The real exchange rate can also be affected by commodity price shocks through their impact on the terms of trade. For instance, an increase in the price of oil improves the international competitiveness of a country which is relatively less dependent on oil. Overall, a lasting deterioration of the terms of trade of a country should result in a depreciation of the real exchange rate of that country. In order to capture such effects, the real price of oil (ROIL) has also been included (ROIL is defined as an index for the oil price deflated by the US producer price index).[18] Since a permanent increase in the price of oil should result in a depreciation of the currency of the country which is relatively more dependent on oil, the expected sign of the variable is *a priori* ambiguous. While the euro area is more oil-dependent than the United States or the United Kingdom, it is less oil dependent than Japan. Nevertheless, given the weight of the former two countries in the trade structure of the euro area, a negative sign appears to be more likely.

- ***Interest rate differentials***

Finally, real interest rate differentials are frequently introduced as an additional determinant of the real exchange rate via the uncovered interest rate parity condition. This relationship provides a link between the short and the medium run. While according to economic theory the interest rate differential should tend to equalise across countries in the long run, the empirical evidence suggests that this is not necessarily the case. Therefore, a less stringent assumption would be that the real interest rate differential is mean-reverting, describing deviations from the exchange rate medium-run equilibrium path. However, since interest rate differentials stay apart over long periods of time, mean-reverting behaviour does not necessarily materialise within the rather short sample span covered in this study. Therefore, even a medium-term impact of the real interest rate differential on the real exchange rate seems to be justifiable in this context.

[18] Amano and van Norden (1998) have studied the relation between the real effective exchange rate of the dollar and the real oil price and found cointegration between them. In their study an increase in the price of oil leads to a real appreciation of the dollar. Dibooğlu (1996) arrives to a similar result when studying the real exchange rates of the lira, the mark and the yen against the dollar, finding a stronger effect for the mark.

Accordingly, the real long-term *(il)* and short-term *(is)* interest rate differential between the euro area and abroad have been included in the analysis. The expected rate of inflation is obtained by smoothing the annualised quarterly change of the consumer price index via an Epanechikov kernel. Assuming that uncovered interest rate parity holds and neglecting risk premia, an increase in the domestic interest rate causes an appreciation of the domestic currency. Long-term real interest rates were proxied by government bond yields since data for discount bonds are not available. As Edison and Melick (1999) have pointed out, a 10-year coupon bond should be a good proxy for a roughly 7-year pure discount bond, which implies that a one percentage point increase of the domestic long-term real interest rate should trigger an increase in the real exchange rate of roughly 7%. However, MacDonald and Nagayasu (2000) show that a consistent implementation of the real interest rate differential needs to take into account the price stickiness of the economy by including an aggregate demand function and a Phillips-curve relationship. If this is included, the sensitivity of the real exchange rate with respect to the real interest rate is also determined by the output gap sensitivity of the price level and by the aggregate demand sensitivity to the real exchange rate, which dampen the maturity effect. Accordingly, the semi-elasticity of the long-term (real) interest rate shocks on the real exchange rate should be below seven.[19]

c. Evolution of the variables

To give an idea of the historical relationships among the economic fundamentals considered in the paper and the real effective exchange rate of the synthetic euro, Chart 1 shows the evolution of the variables. All of them apart from the net foreign asset position and interest rate differentials are in logarithms.[20] The real effective exchange rate of the euro *(q)* declined sharply in the first half of the 1980s, reaching its lowest value in early 1985. Subsequently, it recovered in 1985-86 to moderately oscillate until 1999, when it declined strongly again, getting close to its historic minimum.

Both productivity proxies (*int* and *pro*) followed a similar path up to the mid-eighties, with a relatively large increase during the last years of the seventies. Their diverging evolution from the mid-eighties reinforces, however, the theoretical caveats previously mentioned regarding the similitude between these variables.

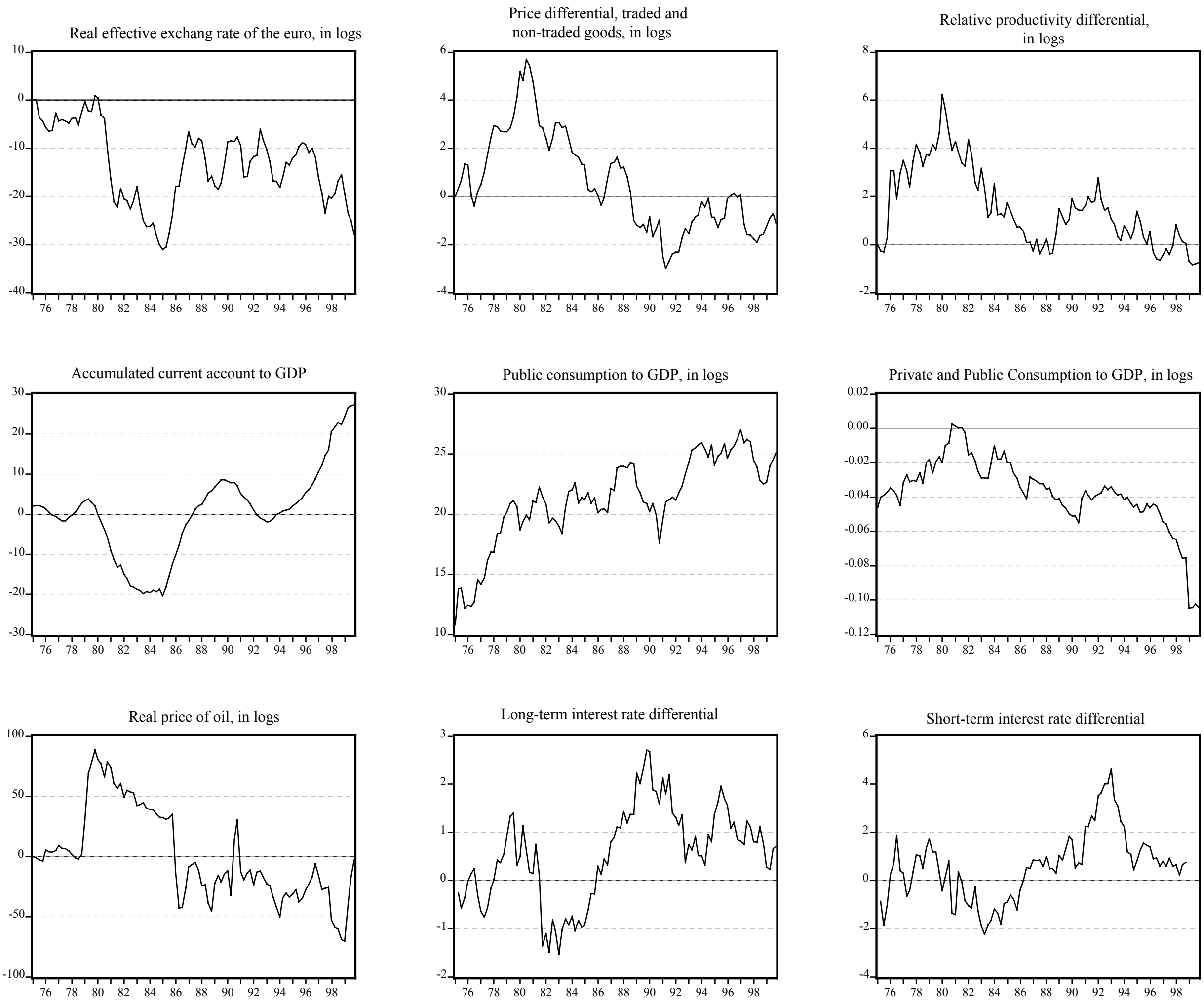
The long-term real interest rate differential *(il)* displays quite some co-movement with the real effective exchange rate. In particular, in the first half of the 1980s the real effective depreciation of the "synthetic" euro coincides with a negative and widening real interest rate differential between the euro area and abroad, which strongly reversed in the following years.

[19] The interest rate differential may be affected by capital controls, which were in place in many of the countries in part of the period under consideration. Using the German long-term interest rate and the US long-term interest rate as representative for the euro area and the rest of the world, respectively, does not have a major effect on these results.

[20] A complete description of the data is provided in the appendix.

In 1995 also, an episode of ("synthetic") euro strength corresponded to a rising real interest rate differential strongly in favour of the euro area. The short-term real interest rates *(is)* move in a broader range than the long-term rates and the effect of the German unification is clearly visible in both.

Chart 1
The real effective exchange rate of the euro
and fundamental factors



Relative public consumption *(gov)* and relative total consumption *(con)* differentials display a very different evolution due to the higher weight of public consumption in the euro area in comparison with its partners. Since the mid-1970s, public consumption to GDP has increased in the euro area relative to its trading partners, suggesting a higher degree of fiscal imbalances in the euro area, which may have weighed on the "synthetic" euro in the long term.

The net foreign asset position (ACA) of the euro area (relative to GDP) shows strong but smooth changes throughout the period, reaching its minimum in 1985, but improving subsequently until the first half of the 1990, when it temporarily levelled off around the time of German unification before surging again.

IV. Econometric methodology

The econometric methodology employed in the paper uses Johansen's cointegration analysis to identify the long-run relationships among the variables. Before estimating the cointegrated vector-error correction by Johansen's method, the stochastic properties of the data are assessed on the basis of a series of unit-root tests. After estimating the long-run relationships, the cointegration parameters are used to perform a permanent-transitory decomposition as suggested by Gonzalo and Granger (1995). The latter transformation yields the PEER (permanent equilibrium exchange rate), while the cointegration analysis allows the construction of the BEER (behavioural equilibrium exchange rate).

a. Unit root tests

The order of integration of the series is assessed using the conventional ADF tests, with lags selected on the basis of information criteria in order to ensure uncorrelated residuals. The ADF test results were confirmed by the tests proposed by Elliot et al. (1996) and Elliot (1999), which have been shown to have superior power properties. More specifically, Elliot et al. (1996) propose two tests, named DF-GLS and P_T, while Elliot (1999) suggests another two, named DF-GLSu and Q_T. If y_t is the series under consideration, then the DF-GLS tests whether ρ=0 in:

$$\Delta y_t^{ld} = \rho y_{t-1}^{ld} + \sum_{i=1}^{k} \Delta y_{t-i}^{ld} + \varepsilon_t \tag{1}$$

where y_t^{ld} is the locally demeaned and detrended series obtained as: $y_t^{ld} = y_t - \hat{\alpha} - \hat{\gamma} t$, and $[\hat{\alpha}, \hat{\gamma}]$ results from regressing y^l on z^l:

$$y^l = [y_1, (1-\bar{\rho}L)y_2, \ldots.., (1-\bar{\rho}L)y_T]$$

$$z^l = [z_1, (1-\bar{\rho}L)z_2, \ldots.., (1-\bar{\rho}L)z_T]$$

where z_t = (1, t)' and $\bar{\rho} = 1 + \bar{c}/T$, and L is the lag operator. The P_T test is defined as $P_T = [S(\bar{\rho}) - \bar{\rho} S(1)] / \hat{\omega}^2$, where S stands for the sum of the squared residual of a regression under the local alternative ($\bar{\rho}$) or under the null, and $\hat{\omega}^2$ is the spectral density obtained from (1). The DF-GLSu and the Q_T are defined in a similar way but the initial observation of the locally detrended series is drawn from the unconditional distribution (i.e., y_1 and z_1 times $(1-\bar{\rho}^2)^{0.5}$).

Asymptotic critical values are provided by the authors for the cases when a constant or a constant and a deterministic trend are included; for the DF-GLS, critical values for the first case correspond to the Dickey-Fuller critical values without deterministic parameters. The authors suggest setting $\bar{c}$ equal to -7.0 for the first case and to -13.5 for the second when using the DF-GLS and the P_T tests, and equal to -10.0 when no deterministic parameters are included, or when using the DF-GLSu and Q_T tests.

For these tests, the lag selection followed a general-to-specific approach, setting the maximum number of lags equal to six (if the coefficient for the highest lag was insignificant at the 90% level, the number of lags was reduced). Deterministic parameters were tested to assess the best model following the testing strategy proposed by Dolado et al. (1990).

b. Vector-error correction models and common trends decomposition

Given the non-stationarity of most variables, cointegration techniques as suggested by Johansen (1995) were employed. This approach enables the detection of long-term equilibrium relationships between the variables.[21] The starting point of this methodology is a vector-error correction model specified as follows:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \mu + \varepsilon_t \tag{2}$$

where y_t is a (n x 1) vector of the n variables of interest, μ is a (n x 1) vector of constants, Γ represents a (n x (k-1)) matrix of short-run coefficients, ε_t denotes a (n x 1) vector of white noise residuals, and Π is a (n x n) coefficient matrix. If the matrix Π has reduced rank ($0 < r < n$), it can be split into a (n x r) matrix of loading coefficients α, and a (n x r) matrix of cointegrating vectors β. The former indicates the importance of the cointegration relationships in the individual equations of the system and of the speed of adjustment to disequilibrium, while the latter represents the long-term equilibrium relationship, so that $\Pi = \alpha\beta'$.

Following Johansen (1995) this standard VEC model can also be expressed in a vector moving average representation, from which further information regarding the key driving variables of the models can be derived:

$$y_t = C\sum_{i=1}^{t} \varepsilon_i + C\eta + C(L)(\varepsilon_t + \eta) \tag{3}$$

where

$$C = \beta_\perp (\alpha_\perp^{'} (I - \sum_{1}^{k-1} \phi_i)\beta_\perp)^{-1} \alpha_\perp^{'} = A\alpha_\perp^{'} \tag{4}$$

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 variables that form the common trends or driving forces of the system. The matrix 'A' represents the loading factors of the common trends, which indicate to what extent each trend influences each variable. Finally, the C-matrix measures the combined effects of the orthogonal components, i.e. the long-run effect of shocks to the system.

[21] In the cointegration analysis we refer to the commonly used terminology of long-term and short-term relationships. This does not contradict the conceptual termini of "current and cyclical" and "medium-term" equilibrium exchange rates introduced earlier.

Closely related to this representation is the decomposition of a non-stationary time series into stationary (T_t) and non-stationary (P_t) elements, as proposed by Gonzalo and Granger (1995):

$$y_t = P_t + T_t .$$

If the time series are cointegrated, the number of non-stationary elements is smaller than the number of series. This implies that the Π matrix has a rank of $r < n$ and if there are n elements in y_t, then there are n-r common factors (f_t). Gonzalo and Granger (1995) identify the two components by imposing that the common factors are linear combinations of the variables in y_t, and that the temporary component does not Granger-cause the permanent component.[22] This is only fulfilled if

$$f_t = \alpha'_{\perp} y_t . \tag{5}$$

This enables Gonzalo and Granger to decompose the individual elements into a permanent and a transitory component, defined respectively as:

$$P_t = \beta_{\perp}(\alpha'_{\perp}\beta_{\perp})^{-1}\alpha'_{\perp} y_t = A_1 f_t \tag{6}$$

$$T_t = \alpha(\beta'\alpha)^{-1}\beta' y_t = A_2 \beta' y_t . \tag{7}$$

V. Econometric results

The econometric strategy followed in the paper has led to the specification of five alternative models, of which four have been retained because it has not been possible to find an encompassing model specification. The discharged model is the only one including the cumulated current account, while the others differ mainly in the definition of productivity and in the treatment of the interest rate differential as a non-stationary or stationary variable.

The models were initially only tested for the period prior to the launch of the euro, for two reasons. Firstly, the launch of the euro itself could mean a structural break. Secondly, it has been claimed that the euro has recently not been in line with economic fundamentals, which could imply that the observations for 1999 and 2000 are outliers. Even in a well-specified model, this could lead to a severe deterioration of the econometric results.[23]

On the basis of each model, a BEER and a PEER have been estimated for the period 1975-98, extrapolated into 2000 and compared with the actual real exchange rate of the synthetic euro. All four models point to the undervaluation of the euro in 2000.

[22] Using the method proposed by Johansen, which derives directly from the C-matrix and the accumulated errors, does not change the results qualitatively (see Hoffmann and MacDonald 2000). The difference between the PEERs obtained from these methods amounts to less than ½%.

[23] In fact, when the models were re-estimated using data up to 1999, only models 2 and 3 proved to be robust, while the results of models 4 and 5 deteriorated.

a. Unit root tests and cointegration findings

The unit root tests were implemented for the 1975:1-1998:4 period. The results – summarised in Table A2 – show that not all series have the same stochastic nature. While *q*, *pro*, *int*, *gov* and *con* are I(1), *is* is I(0), and the results for *il* are ambiguous; it has been treated as I(1) although some Elliot tests indicate that it might be I(0).[24] The results for ACA are somewhat difficult to interpret. While this variable could be considered as I(0) according to some of these statistics, visual inspection of the series casts doubt on this conclusion. Since it displays a rather singular path, it was tested as an I(2) series using the testing strategy proposed by Dickey and Pantula (1987), but this hypothesis was also rejected.[25] Since the series is characterised by very strong inertia, it is most likely an integrated series and therefore has been treated as I(1). The real oil price can be considered either as I(1) or as I(0) with two structural breaks (applying the tests proposed by Lumsdaine and Papell (1997)).

For the VEC model, several specifications have been estimated, with different sets of explanatory variables. In each specification, four centred seasonal dummies were used and the constant was restricted to the cointegrating space. In order to get uncorrelated residuals, the number of lags for each model was selected using Lagrange Multiplier tests for autocorrelation of order one and four. As a first step, the number of cointegrating vectors was decided using the information provided by Johansen's lambda and trace statistics. There has been a growing consensus that both these statistics suffer from a small-sample bias, tending to reject the null hypothesis of no cointegration too often. Two approaches were used to address this problem. The first, proposed by Reimers (1991), adjusts the computed trace or lambda statistics with the factor (*T-nk*)/*T*. The second, suggested by Cheung and Lai (1993), modifies the critical values rather than the tests, using the appropriate response surface regressions. Following the determination of the rank of the Π matrix, the variables were tested for long-run exclusion and weak exogeneity.

Table 2 provides a summary of the results for the five selected models. Results for the variables '*is*' and '*con*' are not reported, as they did not prove to be relevant in any of the specifications. The ***first model*** is the most general specification, comprising five variables (*q*,

[24] According to the real interest rate parity condition, the real interest rate differential should be constant or at least stationary. The finding that the real interest rate differential is I(1) may be due to the relatively short sample or to the existence of a non-stationary risk premium, see Detken et al. (2001).

[25] This strategy consists of testing the null hypothesis of I(2) against I(1) and, if rejected, the null I(1) against the alternative I(0). The starting regression is:

$$\Delta^3 y_t = \rho_1 \Delta^2 y_{t-1} + \rho_2 \Delta y_{t-1} + \sum_{i=1}^{k} \Delta^3 y_{t-i} + \mu + \varepsilon_t$$

Using the critical values of Dickey-Fuller tests, if it is possible to simultaneously reject that ρ_1=0 and ρ_2=0, then the I(2) hypothesis can be rejected. For the present case, the values of the t tests are $t\rho_1$= -2.67 and $t\rho_1$= -2.56, and therefore the null hypothesis is rejected. Another test applied to this series was that proposed by Perron (1997), which allows for a structural break in 1980:4 and rejects that the variable is I(1).

pro, *il*, *ACA*, *roil*). All variables enter significantly in the long-run vector, and the last three are weakly exogenous. While the weak exogeneity of oil prices is fully in line with economic reasoning, this result is rather difficult to explain for the net foreign asset position. In particular, the effects of changes in the oil price should also have an effect on the net foreign asset position. Given the euro area's dependence on oil imports, an increase in the price of oil should lead to a deterioration of the euro area's current account and thus to a decline in its net foreign asset position. Moreover, the coefficient for the accumulated current account suggests that an increase in the net foreign assets position of the euro area actually leads to a euro *depreciation*, whereas an appreciation should be the expected long-term reaction.[26] A further potential problem with this model is that some variables might entail similar information. For instance, the real long-term interest rate may already incorporate the impact of movements in oil prices, insofar as they alters the growth prospects of the economy, and thus the long-term real interest rate itself. Owing to these limitations, this model was not considered further.

Both models 2 and 3 are derived from the first model. After excluding the accumulated current account, the real interest rate differential becomes weakly exogenous and insignificant in the long-run relationship and, therefore, is also excluded from the system. ***Model 2*** shows the main findings for this specification, which includes the real exchange rate, the 'direct' productivity and the real oil price.[27] Both statistics strongly suggest the existence of one cointegration vector. The productivity variable and the real oil price have the expected signs and are significant. The long-run impact of an increase in oil prices seems to be relatively muted, because a 1% increase in the real price of oil causes a decrease in the exchange rate of 'only' 0.2%. However, since this variable tends to show sharp and large movements, its effects can be considerable. Moreover, the real oil price is weakly exogenous, which is a reasonable result, as it is difficult to conceive that the other variables have any influence on it. The productivity differential has a strong impact on the real exchange rate of the euro. Given that this variable may also depend on the relative business cycles, the euro's development appears to relate to the current and expected macroeconomic situation of the euro area relative to its main trading partners. The half-life of a shock is just over one year. This is a rather high speed of convergence to the equilibrium when compared to those obtained by studying the mean reversion of real exchange rates to PPP. However, it is similar to those found in other studies on BEERs (e.g. by Clark and MacDonald (2000) for the US dollar).

[26] Applying the Gonzalo-Granger decomposition, the reason for this behaviour becomes more transparent. The first common trend is mainly associated to *aca* and *q*, and the second and the fourth are also highly related to aca, mainly because their behaviour towards the end of the sample explains the strong upward movement experienced by this variable from the mid-nineties.

[27] In Table 2 and for Model 3, given that ACA in differences appears as an exogenous variable, critical values are shown just for comparison purposes. The same caveat applies to Models 4 and 5.

Table 2. Results of cointegration analysis

	Model 1				Model 2				Model 3				Model 4				Model 5			
	C. vector	Test for exc.	Loading factor	Test for w. exog.	C. vector	Test for exc.	Loading factor	Test for w. exog.	C. vector	Test for exc.	Loading factor	Test for w. exog.	C. vector	Test for exc.	Loading factor	Test for w. exog.	C. vector	Test for exc.	Loading factor	Test for w. exog.
Euro	1.00	10.54	-0.09	16.93	1.00	19.74	-0.15	13.76	1.00	19.73	-0.13	12.86	1.00	14.02	-0.15	7.21	1.00	21.75	-0.14	6.01
Productivity	-12.29	22.85	0.02	14.20	-5.10	9.56	0.03	9.24	-4.19	9.89	0.02	9.15								
Relative prices	-	-	-	-	-	-	-	-	-	-	-	-	-3.01	7.75	0.03	8.57	-1.80	4.83	0.03	7.50
Long-term interest	-7.98	7.36	0.00	0.37	-	-	-	-	-3.99	4.86	0.01	1.05					-6.47	18.84	0.02	3.74
Public consumption	-	-	-	-	-	-	-	-	-	-	-	-	1.74	14.83	-0.03	1.80	1.00	7.30	0.03	2.52
ACA	1.33	12.30	0.00	0.06	-	-	-	-	-	-	-	-								
Oil	0.44	18.21	-0.15	1.13	0.19	8.27	0.08	0.04	-	-	-	-	0.20	10.90	0.05	0.03				
Constant	40.74	24.54	-	-	21.25	17.95	-	-	23.50	17.28	-	-	-20.38	6.48			-2.79	0.22		
Exogenous variables	none				none				aca in differences				il in levels				roil in levels			
Lags	3.00				3.00				3.00				3.00				2.00			
	Lambda		Trace		Lambda		Trace		Lambda		Trace		Lambda		Trace		Lambda		Trace	
	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.	without corr.	with corr.
r = 0	45.93	38.44	81.36	68.09	29.81	26.89	43.19	38.96	25.31	22.83	32.13	28.99	33.62	29.23	63.78	55.46	35.34	32.27	56.29	51.40
r = 1	16.76	14.03	35.43	29.65	9.68	8.73	13.38	12.07	5.58	5.03	6.81	6.14	15.48	13.46	30.16	26.23	11.16	10.19	20.95	19.13
r = 2	13.22	11.06	18.67	15.63	3.70	3.34	3.70	3.34	1.23	1.11	1.23	1.11	12.16	10.57	14.68	12.77	6.97	6.36	9.79	8.94
r = 3	3.78	3.16	5.45	4.56									2.52	2.19	2.52	2.19	2.82	2.57	2.82	2.57
r = 4	1.66	1.39	1.66	1.39																
Critical values	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
According to																				
Cheung and Lai (1993)	43.20	39.66	94.95	89.56	25.03	22.44	39.60	36.28	25.03	22.44	39.60	36.28	33.59	30.43	63.12	58.94	33.59	30.43	63.12	58.94
Osterwald-Lenun (1992)	34.40	31.66	76.07	71.86	22.00	19.77	34.91	32.00	22.00	19.77	34.91	32.00	28.14	25.56	53.12	49.65	28.14	25.56	53.12	49.65

The rationale for ***model 3*** is similar to the one used for Model 2. Starting from the first model, by excluding the real price of oil the collinearity problem is eliminated and the current account can be included as a stationary exogenous variable.[28] The model is similar to the one studied in Makrydakis et al. (2000), but includes an additional exogenous variable. According to the Lambda statistics, the existence of one cointegrating vector is again supported, while the adjusted trace statistic confirms this finding only at the 10% level. Imposing one cointegrating vector, both the speed of adjustment and the coefficient associated to the productivity proxy are similar to those of model 2, while an increase in the interest rate differential by 1 percentage point triggers an appreciation of the euro of 4%. Both variables are significant at standard levels.[29]

The fiscal variable (relative government expenditure) did not prove to be significant in any of the models which include the more direct productivity variable. On the other hand, it becomes significant if this direct variable is substituted by the more indirect productivity variable proxied by the relative price ratio. The specification for ***model 4*** hence includes relative government expenditure, the indirect productivity variable and the real oil price in the long-run relationship, and is similar to the specification studied in Clostermann and Schnatz (2000) for the bilateral euro dollar exchange rate.[30] According to the standard tests, the existence of one cointegration vector can be assumed. The coefficients are significant and have the expected signs. An increase in the relative non-traded to traded goods price ratio generates an appreciation of the euro, while a rise in government consumption leads in the long term to a depreciation of the euro. The real oil price is again weakly exogenous, and has an impact on the real effective exchange rate similar to that found in model 3.

Since the shift of the real interest rate differential to the short-term specification might be controversial following the results of the unit root tests, ***model 5*** reintroduces this variable into the long-run vector and models the lagged real price of oil as an exogenous variable. The tests indicate the presence of one cointegrating vector. Accordingly, while the coefficients of the indirect productivity variable and of the fiscal variable are lower, they are still significant. The real long-term interest rate differential enters with the expected sign and with a coefficient of plausible magnitude, smaller than seven, which is in line with the maturity effect, explained in section III.

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28 Critical values for the trace and lambda statistics are shown only for completeness for models 3, 4 and 5, given the presence of additional exogenous variables, which modifies the asymptotic distribution.

29 The small number of variables of models 2 and 3 allows us testing whether it was correct to introduce the variables in differentials. The results did not change when the euro zone and partners series were introduced independently, so the restriction seems to be in order.

30 For comparison purposes, the long-term interest rate is included a priori as an exogenous variable. Although this variable was found to be I(1), it is still ambiguous whether it is indeed non-stationary. For example, Clostermann and Schnatz (2000) as well as Edison and Melick (1998) find the real long-term interest rate differential to be stationary. For an overview of the properties of real interest rate differentials, see also MacDonald and Nagayasu (2000).

To summarise the cointegration analysis results, the four models that were retained can be classified according to the definition of productivity which is used. For each productivity specification, one model includes the interest rate differential while the other includes the real price of oil as a proxy for the terms of trade, as it was not possible to include these two variables together in the long-run relationship.

b. Decomposition into permanent and transitory components

Table 3 reports the orthogonal complements, the loading factors of the common trends, and the impact matrix for the moving average representation of models 2 to 5, computed from equation 4 (in the Appendix, Table A3 shows the results for model 1). These matrices provide some additional information, as well as some further insight regarding the internal consistency of the exercise.

For each model, the first matrix shows the alpha orthogonal components, indicating which variables drive the common trends.[31] Each trend is mainly driven by the variable with the highest absolute value, which is highlighted in bold in the table. In the current exercise, one would expect the common trends to be driven by the fundamental determinants of the real exchange rate rather than by the real exchange rate itself. The second matrix illustrates which variables are most affected by the common trends, i.e. how the impact of the common trends is distributed amongst the variables, and which common trend affects the real exchange rate the most (the corresponding coefficient is in bold in the table). The third matrix is the main coefficient in the moving average representation of the standard VEC model, which is the product of the two aforementioned matrices and combines the information given therein (this is the so-called C matrix given by $C=A\alpha_{\perp}'$ and the numbers in bold in the table indicate the coefficients which are statistically significant).

This impact matrix discloses more information and allows the model to be checked for consistency. It shows which variables exert a cumulative impact on the real exchange rate, it indicates the exogeneity properties of the variables, and it provides information regarding the persistency of individual time series. In particular, the individual cumulative shocks should have a significant (and correctly signed) impact on the real exchange rate, as indicated by the first column of the C-matrix.[32] Weak exogeneity is indicated if none of the cumulative shocks are significant for a variable besides shocks to itself. The diagonal of the matrix indicates the inertia of the system.

[31] These results should be interpreted cautiously, however, since the statistical significance of the orthogonal elements cannot be computed.

[32] The t-ratios in brackets are based on the asymptotic standard errors suggested by Paruolo (1997).

Table 3: The Orthogonal Components and the Long-run Impact Matrices

MODEL 2	q	pro	roil
$\alpha_{\perp}^{1}$	0.000	0.000	**1.000**
$\alpha_{\perp}^{2}$	0.180	**1.000**	0.000
A1	-0.110	0.016	0.985
A2	**3.234**	0.767	4.194
$\Sigma\varepsilon_{q}$	**0.582**	**0.138**	0.755
	[2.17]	[2.24]	[0.54]
$\Sigma\varepsilon_{pro}$	**3.232**	**0.767**	4.194
	[4.78]	[4.94]	[1.18]
$\Sigma\varepsilon_{roil}$	**-0.110**	**0.016**	**0.985**
	[-3.23]	[2.09]	[5.49]

MODEL 3	q	pro	il
$\alpha_{\perp}^{1}$	0.000	0.000	**1.000**
$\alpha_{\perp}^{2}$	1.000	**5.298**	0.000
A1	**1.873**	-0.209	0.815
A2	0.599	0.143	0.013
$\Sigma\varepsilon_{q}$	**0.599**	**0.143**	0.013
	[2.28]	[2.39]	[0.38]
$\Sigma\varepsilon_{pro}$	**3.175**	**0.755**	0.071
	[5.22]	[5.48]	[0.86]
$\Sigma\varepsilon_{il}$	**1.873**	-0.209	**0.815**
	[2.17]	[-1.07]	[6.95]

MODEL 4	q	int	gov	roil
$\alpha_{\perp}^{1}$	1.000	**5.426**	0.000	0.000
$\alpha_{\perp}^{2}$	0.000	0.000	0.000	**1.000**
$\alpha_{\perp}^{3}$	0.000	0.000	**1.000**	0.000
A_1	0.450	-0.208	0.079	0.257
A_2	-0.189	-0.001	0.015	0.900
A_3	**-2.264**	-0.149	1.425	3.436
$\Sigma\varepsilon_{q}$	0.450	**0.208**	0.079	-0.257
	[1.53]	[3.00]	[0.67]	[-0.25]
$\Sigma\varepsilon_{int}$	2.444	**1.128**	0.427	1.393
	[1.73]	[3.38]	[0.75]	[0.28]
$\Sigma\varepsilon_{gov}$	**-2.264**	0.149	**1.425**	3.436
	[-2.82]	[0.79]	[4.45]	[1.25]
$\Sigma\varepsilon_{roil}$	**-0.189**	-0.001	0.015	**0.900**
	[-3.29]	[-0.09]	[0.67]	[4.57]

MODEL 5	q	int	gov	il
$\alpha_{\perp}^{1}$	1.000	**4.551**	0.000	-0.920
$\alpha_{\perp}^{2}$	0.000	0.000	**1.000**	0.000
$\alpha_{\perp}^{3}$	0.869	1.000	0.000	**5.894**
A_1	0.067	0.242	0.045	-0.054
A_2	**-0.837**	0.205	1.116	0.004
A_3	0.666	-0.023	0.024	0.118
$\Sigma\varepsilon_{q}$	**0.646**	**0.222**	0.067	0.049
	[2.66]	[2.71]	[0.65]	[1.61]
$\Sigma\varepsilon_{int}$	0.972	**1.079**	0.231	-0.125
	[1.39]	[4.57]	[0.79]	[-1.43]
$\Sigma\varepsilon_{gov}$	**-0.837**	0.205	**1.116**	-0.004
	[-2.35]	[1.70]	[7.44]	[-0.10]
$\Sigma\varepsilon_{il}$	**3.866**	-0.358	-0.102	**0.746**
	[4.08]	[-1.12]	[-0.26]	[6.27]

Model 2 fulfils all expected requirements (see Table 3). The matrix of the orthogonal complement of alpha confirms that one common trend of model 2 is associated to the productivity differential, while the other is explained by the real price of oil. The real exchange rate is mostly driven by the common trend corresponding to the productivity differential, as indicated by the loading matrix. The impact matrix (C-matrix) shows that there is still a considerable inertia in q, a feature which holds for all four models and reflects to some extent the non-stationarity of the real exchange rate. In addition, the first column of the C matrix shows that the productivity differential is significantly and positively related to the real exchange rate, while increasing oil prices exert significant downward pressure on the real exchange rate. The productivity differential is positively related to the other two: this could be explained by the possibility that a real appreciation of the euro increases international competition and fosters the productivity of companies in the euro area relative to their competitors in the partner countries; increases in oil price have a similar effect. The last column of the C matrix supports the notion that the real oil price is a weakly exogenous variable since it is only driven by itself.

In model 3, the common trends are associated with the productivity differential and with the real interest rate differential. Both common trends drive the real exchange rate, but the common trend related to the real interest rate differential appears to have a stronger impact. In the first column of the C matrix, the positive and significant impact of the common trends associated with the productivity differential and with the real interest rate differential is reinforced. As in the previous model, the productivity differential is positively influenced by the real exchange rate, while the long-term real interest rate differential has no statistically significant impact on the productivity differential. Furthermore, the hypothesis that the real interest rate differential is weakly exogenous is supported by the third column of the C matrix, as the other two variables are insignificant.

For model 4, one common trend is linked to the relative government expenditure ratio, while the other two relate to the relative price ratio of non-traded to traded goods and to the real price of oil. The real exchange rate is mainly driven by the common trend corresponding to the relative government expenditure as a ratio to GDP, while the common trend related to the real price of oil mainly explains the oil price itself. Finally, the common trend corresponding to the relative price ratio is also related to the oil price variable. The impact matrix confirms a significant impact of the fiscal variable and of the oil price on the real exchange rate, but fails to support an impact of the indirect productivity variable on the real exchange rate. A positive shock to the real exchange rate (appreciation) translates into an increase of the relative price ratio. Neither the relative government expenditure ratio nor the oil price is significantly related to any of the other variables, supporting their characterisation as weakly exogenous variables.

In model 5, the orthogonal complements to alpha show that each common trend is associated to one variable other than the real exchange rate. The real exchange rate appears to be driven mainly by the common trend linked to the government expenditure differential, which has the largest loading factor. Again, relative government consumption and the real interest rate differential have both the expected sign and a significant impact on the real exchange rate, while the relative price ratio has the correct sign but is insignificant. Moreover, there is support for the weak exogeneity of relative government expenditure as well as of the interest rate differential. In both models 4 and 5, shocks to the real exchange rate have a positive and significant impact on the relative price differential. This is a reasonable result, as it implies that the relative effect of a real appreciation is higher on the wholesale price index than on the consumer price index, as the former includes more sectors open to international competition. However, both models seem to have one drawback, as they do not closely arrive at the same conclusions obtained using the VEC specification. In particular, the relative price ratio is insignificant in the C-matrix, although all the variables are signed correctly (however, the former may be related to the asymptotic properties of the standard errors).

In summary, the permanent-transitory decomposition indicates that the main forces driving the real exchange rate are productivity and either the real interest rate differentials or the real price of oil. This is apparent both from the tests for weak exogeneity and from the structure of the common trends loading matrices.[33]

c. Estimated permanent and behavioural equilibrium exchange rates.

This section examines the BEER and PEER derived from models 2 to 5. Table 4 shows the correlation between the individual BEERs in the upper right triangle and correlation between the PEER models in the lower left triangle. The diagonal displays the correlation between the BEER and the PEER projections for each model. All correlations are positive, and they are much higher for the PEER than for the BEER projections. Moreover, for each model both projections have a coefficient larger than 0.8, which means that there is a sizeable degree of co-movement between them.[34] This suggests that a similar amount of transitory shocks are filtered out in each model, and that some variables of the various models might indeed capture a comparable impact on the real exchange rate. For instance, the PEER specifications with the real interest rate differentials and those with the oil prices might become more similar, since higher oil prices might be related to relatively decreasing long-term real interest rates in the euro area. Therefore, these variables can have a more comparable permanent impact on the exchange rate after filtering out the different transitory disturbances. Overall, a somewhat

[33] The common trends confirm the close relation between the real exchange rate and the productivity variables, measured both directly and indirectly. In model 4, the main driving force of the system is the real price of oil. The second common trend from model 5 basically explains the relative price differential, but also plays a role in the behaviour of the euro up to the second half of the eighties.

[34] Actually this should be the case by definition and serves rather as a consistency check.

higher correlation is found between models 3 and 5, which reflects the specification based on the real interest rate and two proxies for productivity differentials, while model 4 is generally the least correlated with the other models.

Table 4: Correlation of the Forecast Series

	BEER2	BEER3	BEER4	BEER5	
PEER2	0.847	0.615	0.540	0.376	BEER2
PEER3	0.901	0.877	0.188	0.783	BEER3
PEER4	0.869	0.766	0.840	0.335	BEER4
PEER5	0.861	0.948	0.828	0.886	BEER5
	PEER2	PEER3	PEER4	PEER5	

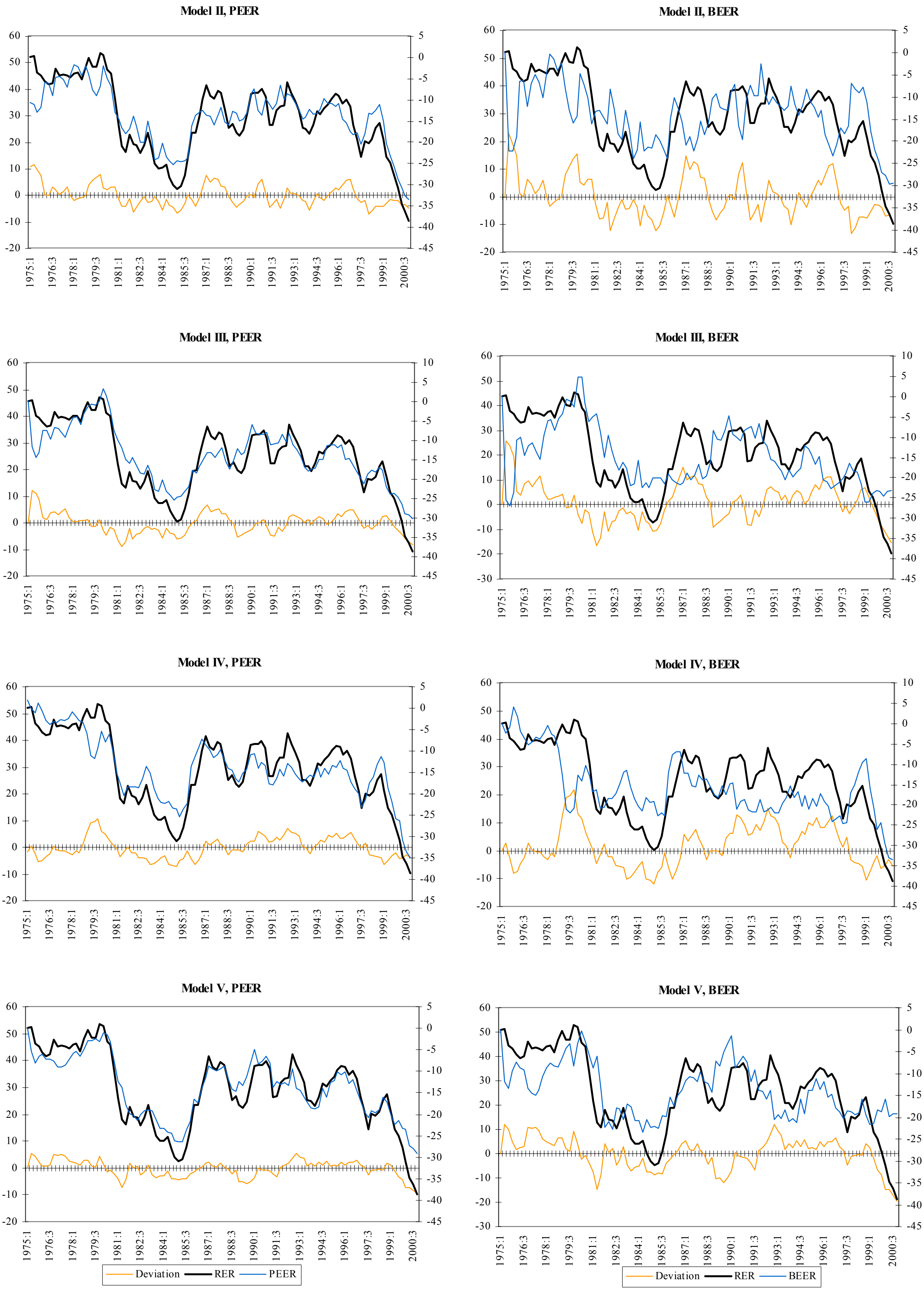
Chart 2 shows for each model under consideration the actual real effective real exchange rate of the euro (dark line) with its fitted value using the long-run coefficients (BEER) and with its permanent component derived from the common factor representation (PEER).[35] The series oscillating around the zero-line is the difference between the real effective exchange rate and the BEER or PEER respectively. For the four models, the estimated parameters using the sample period up to 1998:4 were employed to compute projected values up to 2000:4 (using the actual values of the series).[36]

As expected, the PEER is less volatile than the BEER and, as documented by the correlations, the differences between the actual series and its fitted and permanent values is neither large nor persistent. The models often give conflicting signs, not only with regard to the magnitude, but also with respect to the direction of deviation of the exchange rate from the computed equilibria. However, there are some periods in which all models point to the same direction of misalignment. While the euro was rather close to its fundamental value in the seventies and the first half of the nineties, all the models show an undervaluation of the euro in the mid-1980s, which coincided with the dollar’s strength prior to the Plaza agreement. The subsequent overreaction of the exchange rate in the opposite direction is also well documented, as the models unanimously show some overvaluation of the euro in 1987. In the 1990s, strong signs of some overvaluation of the euro become evident in 1995-1997, which again corresponds to a period of weakness of the US dollar against major European currencies. After 1997, the models give very different signals depending on the specification of the model, and in particular, on whether the oil price is part of the specification.

35 The Gonzalo-Granger decomposition of the series was performed using demeaned data. Therefore, as the models were estimated with the constant restricted to the cointegrating vector, the mean of the transitory component is null. However, in order to plot the PEER together with the BEER and the actual series of the euro, the mean of the series was added to the estimated PEER, rather than decomposing it between the permanent and the transitory component.

36 In fact, the in-sample forecasts (up to 1998) and the out-of-sample forecasts are not strictly comparable. The estimation over the full period would probably reduce the magnitude of the undervaluation found (i. e. the results presented below could be biased to some extent).

Chart 2: Real effective exchange rate (RER) of the euro, PEERs and BEERs



Focusing on the last three years, Table 5 shows the difference between the real exchange rate and its permanent component (column headed by PEER), or its long-run value (column headed by BEER) according to the models.

Table 5: Forecast differences

	Model 2		Model 3		Model 4		Model 5	
	PEER	BEER	PEER	BEER	PEER	BEER	PEER	BEER
1998:1	--	---	-	-	--	-	.	.
1998:2	--	---	.	.	--	-	.	.
1998:3	--	--	.	.	--	-	.	.
1998:4	-	--	+	++	--	-	.	+
1999:1	-	--	+	++	---	--	.	+
1999:2	-	--	.	.	---	--	.	-
1999:3	.	-	.	.	---	-	-	--
1999:4	.	-	-	--	---	-	-	--
2000:1	.	-	--	--	--	--	--	---
2000:2	-	--	--	---	--	--	--	---
2000:3	-	-	--	---	--	-	--	---
2000:4	-	--	--	---	---	--	--	---

+/++/+++ indicate overvaluation > 2%, 5%, 10%, -/--/--- indicate undervaluation > 2%, 5%, 10%, . indicates deviation < 2%.

At the beginning of 1998 the euro was undervalued according to the BEERs and PEERs, and this undervaluation was reduced or even over-corrected by the end of the year in three of the four models, while it built up further in the case of model 4. For the period around the launch of the euro, the results are ambiguous. Owing to the decline in oil prices which took place until 1999, models 2 and 4 show that the euro was undervalued in effective terms in the quarter after its launch. The results change if the specifications include the real interest rate differential instead of the oil price (models 3 and 5): according to model 5, the euro was by and large in equilibrium at the beginning of 1999, while model 3 shows some overvaluation at that time.

Beginning from the second half of 1999, all the models point in the same direction, indicating that the euro was again undervalued. Also in this case, the amount of undervaluation is smaller in the specifications which take into account the surge in oil prices since the first quarter of 1999. The sharp increase in the real price of oil narrowed considerably the amount of undervaluation disclosed by these models for the first quarter of 1999. Models 3 and 5, on the contrary, suggest a widening of the undervaluation of the euro towards the end of 1999. In any case, the deviations from the fundamentals were not extraordinarily large by historical standards at that time. During the year 2000, however, all the models unanimously describe the euro as departing further from its equilibrium, pointing to an increasingly undervaluation in the course of 2000. According to model 3, in the last quarter of 2000 the euro's undervaluation was at its highest level since the mid-eighties, while model 5 even points to the highest undervaluation in the entire sample period. By contrast, the undervaluation of the euro was not unprecedented in terms of models 2 and 4, since the slump in oil prices until 1999 would have required an equilibrium appreciation of the euro, which did not materialise. On the contrary, the subsequent surge in oil prices shifted the equilibrium exchange rate downwards, narrowing the amount of undervaluation. However, the fact that all models pointed to an undervaluation of a magnitude not observed since the mid-eighties supports the viewpoint that the euro went out of line with fundamentals in the course of 2000.

The estimates of the disequilibrium at any point in time are surrounded by a large degree of uncertainty, which adds to the difficulty to derive a precise equilibrium level of the effective exchange rate. From the BEER and the PEER estimates, the computation of standard error bands is not straightforward since both are non-stationary. Accordingly, standard errors have been computed for the BEER disequilibrium term, which is stationary and, hence, has a finite variance. In this case, the standard error of the disequilibrium is estimated by the square root of $\beta' R_1' R_1 \beta$, where R_1 is the residual from the regression of Y_{t-1} on all unrestricted variables in the VECM.[37] The estimated standard errors for the models estimated in this paper range from above 4 in model 5 to just above 6 in model 3.

Overall, it is a rather exceptional event when at least three of the four models point to an over- or undervaluation of the euro, as illustrated in Chart 3. The chart shows the deviations of the exchange rate from the four equilibrium rates (in percentage points). The dark bars indicate periods in which three of the four models indicate a misalignment of the euro, which is defined as realisations lying outside the two standard deviations band. The light-grey bars are assigned to periods in which at least three of the four models moved outside 1.5 standard deviation bands. The upper panel shows periods when the models signal an euro overvaluation, while the lower panel points to episodes in which the models signify an euro undervaluation.

Some sporadic periods of euro overvaluation are found in the late seventies and in 1987, in the aftermath of the “Louvre Accord”. There are also some more protracted periods of overvaluation in the 1990s, usually reflecting episodes of dollar weakness. The lower panel reveals a cluster of joint signals pointing to a protracted euro undervaluation in the first half of the 1980s, coinciding with the period of extraordinary dollar strength prior to the “Plaza Accord”. However, 2000 contains the only period since the mid-eighties in which all the models again jointly signal an undervaluation of the euro, and it is the only episode, in which the estimates of three of the four models lie outside the 95% confidence band. This result should be interpreted with some caution, however, since it refers to the only episode where the estimates have been employed out-of-sample, which could potentially introduce a bias into the results.

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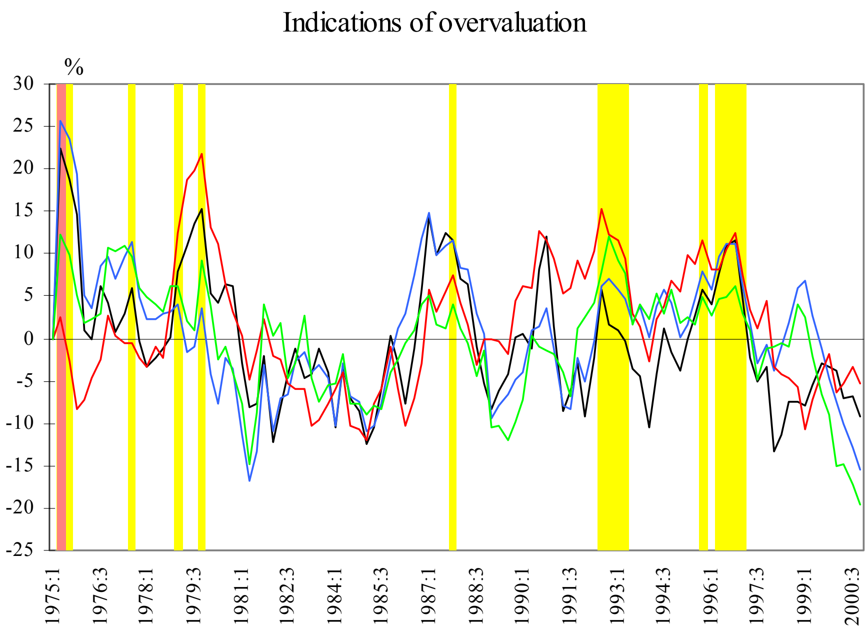
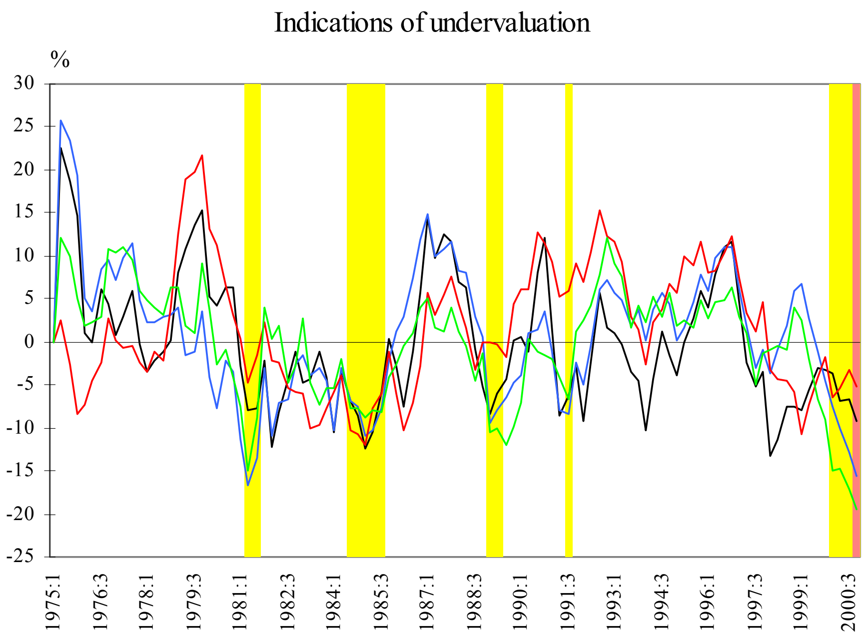
[37] See Johansen (1995) for a discussion of the distribution of $\beta' X_t$.

Chart 3. Common signals of euro over-/undervaluation





Note: The lines show the deviations of the each model from the BEER. The grey bars show periods in which at least three of the four models indicate an overvaluation (upper panel) or an undervaluation (lower panel). The light-grey (dark-grey) bars assign periods in which at least three of the four models are outside 1.5 (2) standard deviation band.

VI. Conclusions

This paper is based on a very comprehensive and consistent data set for the euro area and its most important trading partners, which has been compiled for the period from 1975 to the present on a quarterly basis. A synthetic real effective exchange rate of the euro against the twelve main trade partners of the euro area, as well as fundamental variables in effective terms, have been constructed. The selection of the main determinants of the euro effective exchange rate follows the most relevant theoretical models. These take into account the effects of (sectoral) productivity differentials, real interest rate differentials, government expenditures, time preferences, net foreign assets evolution and the terms of trade (using the real price of oil as a proxy).

Applying Johansen's procedure, four models (models 2 to 5 in the text) were estimated and evaluated using the sample size up to the fourth quarter of 1998. These models imply that the euro is affected by the productivity differential, measured directly or indirectly. Two of the models also include the real price of oil, while the others include the real long-term interest rate differential. In the two models where the productivity differential was measured indirectly by the relative price variable, relative government spending was also significant. In contrast, the inclusion of the net foreign asset position did not lead to a consistent model, and the proxy for the time preference rate and the short-term interest rate differential proved to be insignificant. Overall, the euro appears to be mainly affected by productivity developments, real interest rate differentials, and external shocks due to the oil dependence of the euro area.

The Behavioural Effective Exchange Rates (BEERs) obtained from the cointegrating vectors were compared with the Permanent Effective Exchange Rates (PEERs) obtained using the Gonzalo-Granger decomposition. These measures of equilibrium exchange rates are rather similar, the PEERs being smoother than the BEERs. Both indicate that the euro was close to its fundamental value (or slightly overvalued) in the seventies and in the first half of the nineties; equally, they detect a sizeable undervaluation in the first half of the eighties. During the first year after its launch, the euro experienced a strong depreciation in effective terms. While the estimated models account for some of that depreciation, they also unanimously indicate undervaluation by the end of 1999. In the course of 2000, all four models suggest that the euro deviated further from equilibrium, which supports the judgement that the euro was undervalued in effective terms towards the end of 2000.

The analysis in this paper shows that it is far from easy to estimate the equilibrium exchange rate precisely and to come up with one agreed relationship, as indicated by the failure to find a unique model encompassing all four specifications analysed. This difficulty gives an idea of how complicated it can be to derive an exact quantification of the amount of misalignment of a currency. However, observing that periods when the majority of these models pointed to some significant misalignment in the same direction were rather infrequent supports the conclusion that the euro went out of line with economic fundamentals in 2000.

APPENDIX.

Data description

The Euro data comprises France (FR), Spain (ES), Germany (DE), Greece (GR), Austria (AT), Italy (IT), Portugal (PT), Belgium (BE), Netherlands (NL), Ireland (IE), Finland (FI). The partners countries are: Australia (AU), Canada (CA), Switzerland (CH), Denmark (DK), Great Britain (GB), Hong Kong (HK), Japan (JP), South Korea (KR), Norway (NO), Sweden (SE), Singapore (SG), United States (US).

Quarterly data (1974:1-1999:4), except for some series where data was converted from annual source applying a spline method. Data for 2000 was partly estimated.

The weights used to compute the variables are average overall trade weights (taking into account third market effects) for 1995-96-97. These weights are used to compute EMU variables and partners variables (unless otherwise stated, aggregation is done by geometric average). In rare cases, when data were missing for a certain country, the weights were re-based skipping that country.

Nominal exchange rate: - IFS, line rf (unit per dollar)

Consumer Price Index: - IFS, line 64. For EMU countries, harmonised CPI from 1999:1 onwards (line 64h). For HK, national sources up to 1994:4 : from 1980:4 to 1994:4 composite CPI (it covers 90% of household); from 1974:3 to 1983:3 a index is constructed by averaging three CPI (each one covering different households according to their monthly expenditure, their weights are given by the percentage of households belonging to a given interval of monthly expenditure).

Wholesale Price Index: - IFS, line 63 for all countries except PT, where OECD data is used (line ppiamp01.ixob). For HK, data starts in 1993:1, for PT in 1990:1. For some countries missing data were obtained from paper version of IFS.

Oil price: -IFS, spot price index (line 00176AADZF)

Long-term interest rates: - IFS, line 61. No data for HK and SG. For SE, from 1987:1 onwards, OECD data (line irltgv02). For ES, no data prior to 1978:2. For FI, OECD data from 1993:1 onwards; from 1991:4 to 1992:4, BIS data, secondary market for government bonds (ten years), (line BISHGBAFI03); from 1982:2 to 1991:3, BIS data, secondary market for government bonds (three to seven years) (line BISHGAAFI22); from 1980:4 to 1982:1, BIS data, secondary market public issues (four-five years), (line BISHHLAFI94); from 1974:1 to 1980:3, BIS data, interest rates for deposits at two years, (line BISPHAFI94). IE, IFS up to 1998:4, OECD afterwards. PT, no data up to 1976:1.

Short-term interest rates: -IFS, line 60b (money market rate). For CA, CH: OECD, line irt3. For GB, IFS line 60c (treasury bill rate). HK, national sources from 1980:1 to 1993:3 (converted), IFS line 60b onwards. KR, line 60 (discount rate, end of period) from 1974:1-1976:3, line 60b onwards. FI, line 60 up to 1977:4, line 60b onwards. GR, line 60l (commercial banks deposits of three to twelve months) up to 1985:2, line 60c onwards. IE, line 60b except for a few missing observations where line 60c was used. PT, line 60 up to 1980:4, line 60b onwards.

Net Foreign Asset Position. It is the sum of accumulated current account position up to one period divided by the current GDP, both measured in dollars (GDP corresponds to the GDP of each quarter, no annualised).

GDP is obtained from IFS, line 99b (for some countries, annual data converted into quarterly).

For NFA, 1973 is used as the starting date, employing Milesi-Ferretti database. Current account data from IFS, line 78aldzf (for some countries, missing data was covered using paper version of IFS or the M-F database).

Productivity: Constant GDP divided by employment. Constant GDP is obtained from IFS, line 99bv (in most cases, annual data converted to quarterly). Employment from OECD (line emestt02) or AMECO. In the case of Germany, employment was obtained from the BIS data base and completed with data from the OECD. To account for the effect of German unification, a regression using German productivity was estimated introducing an impulse dummy variable (value one in 1991:1 and zero otherwise).

Government consumption: For EMU countries, data from AMECO (code 1.0.0.0.UCTG), except Germany (data from IFS). For partners countries, data from IFS. GDP at current prices from the same source as government consumption.

Total consumption: Private consumption (line 96F), government consumption (line 91F) and GDP at current prices from IFS.

Methodological changes from ESA79 to ESA95 generate breaks in the series of government consumption, and to a lesser extent to private consumption, for some EMU countries. While there are no appreciable changes when computing total consumption, when computing government consumption as a fraction of GDP it is necessary to use series that take into account those methodological changes.

Table A1. Weighting scheme for the construction of the variables

	weight w_j		weight g_i
Austria	2.89	Australia	1.13
Belgium	7.98	Canada	1.96
Finland	3.27	Denmark	3.50
France	17.75	Hong Kong	3.90
Germany	34.49	Japan	15.01
Greece	0.736	Korea	4.91
Ireland	3.76	Norway	1.70
Italy	13.99	Singapore	3.50
Netherlands	9.16	Sweden	6.23
Portugal	1.07	Switzerland	8.84
Spain	4.90	United Kingdom	24.26
		United States	25.05
	100.00		100.00

Table A2. Results of unit root tests

	Model with constant and trend					Model with constant					Model without constant	
	ADF	DFGLS	PT	DFGLSU	QT	ADF	DFGLS	PT	DFGLSU	QT	ADF	DFGLS
Variables in levels												
q	-2.64	-1.53	15.06	-1.70	6.88	-2.70*	-0.80	12.92	-1.80	8.36	-0.99	-0.99
Pro	-2.45	-1.76	19.00	-2.12	6.78	-1.54	-1.64***	6.60	-1.64	9.28	-1.47	-1.55
Int	-2.64	-1.90	13.95	-2.34	5.92	-1.42	-1.17	6.27	-1.19	11.84	-1.46	-1.16
Il	-2.30	-2.38	5.08**	-2.38	2.81**	-1.95	-1.56	4.08***	-1.91	5.35***	-1.59	-1.50
Is	-2.49	-2.58***	7.60	-2.60	4.16	-2.52	-1.82***	4.67	-2.37	5.81***	-2.12**	-2.05**
Gov	-2.11	-1.58	19.52	-2.49	8.02	-2.16	-0.36	46.07	-2.61***	25.06	0.60	-
Con	-1.99	-1.36	15.53	-2.05	5.91	-0.32	-1.01	6.06	-0.92	9.87	0.56	0.42
ACA	-2.33	-2.08	4.95**	-2.36	2.49***	-1.60	-1.73***	1.86*	-1.73	3.15***	-1.71*	-1.71***
Roil	-2.79	-1.75	8.98	-2.24	3.40***	-1.28	-1.52	3.94***	-2.09	3.58***	-1.33	-0.68
Variables in first differences												
Q	-5.77***	-6.97***	2.10***	-6.98***	1.15***	-5.79***	-5.26***	0.15***	-5.28***	0.32***	-5.82***	-4.54***
Pro	-4.28***	-3.86***	5.98**	-3.88***	3.22*	-4.44***	-8.49***	0.32***	-8.47***	0.65***	-4.44***	-6.06***
Int	-5.63***	-5.21***	0.02***	-5.20***	0.01***	-5.64***	-5.17***	0.01***	-5.17***	0.02***	-5.66***	-4.58***
Il	-3.82**	-3.62***	7.74	-3.90***	3.40*	-3.84***	-2.65***	3.72*	-3.95***	3.32***	-3.87***	-6.91***
Is	-7.49***	-8.37***	2.15***	-9.68***	1.15***	-7.51***	-6.22***	0.81***	-9.92***	1.20***	-7.54***	-7.96***
Gov	-6.51***	-1.41*	34.24	-2.93*	8.86	-6.41***	-0.36	40.14	-5.01***	11.58	-6.37***	-12.70***
Con	-7.18***	-1.68	33.36	-2.98*	8.62	-6.86***	-0.43	89.46	-2.98***	20.84	-6.83***	-14.60***
ACA	-2.69	-2.36	7.85	-2.59	3.01*	-2.56	-2.02**	4.04*	-2.59**	2.95***	-2.50**	-2.50***
Roil	-5.34***	-1.69	45.60	-2.36	10.68	-5.24***	-0.33	54.50	-2.79**	7.88	-5.23***	-3.58***

Note: Three/two/one asterisks for rejection at 2.5%/ 5%/ 10%. The model with a constant is not implemented if the trend was significant, the model without deterministic parameters is not used if the constant was significant.

Table A3. The orthogonal complements and the long-run impact matrix

MODEL 1	q	pro	il	aca	roil
$\alpha_{\perp}^{1}$	-0.455	0.111	-0.010	0.829	0.306
$\alpha_{\perp}^{2}$	0.001	-0.197	-0.980	0.019	-0.009
$\alpha_{\perp}^{3}$	-0.152	-0.967	0.195	0.059	-0.028
$\alpha_{\perp}^{4}$	0.706	-0.059	0.027	0.556	-0.435
A1	2.182	0.345	0.409	5.072	-3.164
A2	-5.193	-0.045	-1.107	-4.445	3.836
A3	-1.089	-0.298	0.496	4.573	-10.569
A4	2.054	0.231	0.293	3.624	-3.799
$\Sigma\varepsilon_{q}$	0.615 [2.13]	0.051 [1.53]	-0.056 [-1.18]	-0.449 [-1.00]	0.366 [0.33]
$\Sigma\varepsilon_{pro}$	2.194 [1.42]	0.322 [1.80]	-0.234 [-0.93]	-3.199 [-1.32]	9.337 [1.58]
$\Sigma\varepsilon_{il}$	4.913 [2.42]	-0.011 [-0.05]	1.186 [3.58]	5.298 [1.67]	-5.895 [0.76]
$\Sigma\varepsilon_{aca}$	2.790 [1.45]	0.396 [1.78]	0.511 [1.63]	6.404 [2.13]	-5.283 [-0.72]
$\Sigma\varepsilon_{roil}$	-0.146 [-2.66]	0.014 [2.24]	-0.006 [-0.68]	-0.111 [-1.30]	0.947 [4.54]

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