The determinants of the euro-dollar exchange rate

Synthetic fundamentals and a non-existing currency

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Summary

At the beginning of 1999 the euro was launched as a common currency in 11 European countries. This paper addresses empirically the medium to long-term forces driving the real euro-dollar exchange rate. Constructing a synthetic euro-dollar exchange rate over a period from 1975 to 1998 and applying cointegration approaches, four factors are identified as fundamental determinants of the real euro-dollar exchange rate: the international real interest rate differential, relative prices in the traded and non-traded goods sectors, the real oil price and the relative fiscal position. A single equation error correction model outperforms multivariate models and seems to be best suited to analyse and forecast the behaviour of the euro-dollar exchange rate in the medium-term perspective. If this model is applied to the current developments in foreign exchange markets, the external value of the euro appears to be rather low in the winter of 1999/2000.

JEL F31, F47

Key words: real exchange rates, fundamentals, cointegration, forecast.

Zusammenfassung

Zum Jahresbeginn 1999 wurde der Euro als gemeinsame Währung in 11 europäischen Staaten eingeführt. In der vorliegenden Studie werden die mittel- bis langfristigen Determinanten des Euro empirisch untersucht. Unter Verwendung eines synthetisch berechneten Euro/Dollar-Wechselkurses werden auf der Basis der Kointegrationsanalyse vier Faktoren als fundamentale Bestimmungsgründe des realen Euro/Dollar Wechselkurses identifiziert: die internationale Realzinsdifferenz, das relative Preisverhältnis gehandelter und nicht-gehandelter Güter, der reale Ölpreis und die relative Staatsausgabenquote. Es zeigt sich, daß ein Eingleichungsfehlerkorrekturansatz zu besseren Ergebnissen führt als ein Vektorfehlerkorrekturmodell und damit am besten geeignet scheint, um das Verhalten des Euro/Dollar-Wechselkurses über die mittlere Frist zu analysieren und zu prognostizieren. Eine Anwendung des Modells auf die derzeitige Wechselkurssituation legt die Vermutung nahe, daß der Außenwert des Euro im Winter 1999/2000 recht niedrig bewertet ist.

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1. Introduction

When considering the factors determining exchange rate movements a familiar starting point is the theory of (relative) purchasing power parity (PPP), according to which the inflation differential at home and abroad is reflected in a corresponding change in the nominal exchange rate. If the PPP were actually the only driving force behind the fluctuations in the nominal exchange rate in the medium term, the real exchange rate would have to be a constant over this time horizon. Econometrically, however, it has proved difficult to confirm this relationship convincingly, when the period of more flexible exchange rates since the mid-seventies is reviewed.[1]

In the light of the well-known weaknesses of these tests, various studies have recently been carried out to improve the empirical substantiation of the PPP hypothesis, aiming to increase the power of the test procedures. Firstly, the length of the estimation period was considerably extended in some studies.[2] Secondly, the flexibility of the test method was improved by taking fractional cointegration into account.[3] Thirdly, the number of observations was increased by using panel data.[4] Many of these studies found some evidence for the validity of the PPP hypothesis, at least for the very long term. However, the half-life periods derived from these studies as a measure of the speed with which shock-related deviations of the real exchange rate are restored are decidedly long at approximately four to five years.

A number of studies, however, went beyond this simple, essentially univariate approach by adding further explanatory variables to exchange rate equations. In contrast to the (univariate) PPP analyses, which implicitly assume a constant equilibrium exchange rate,

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1 See, for example, Froot and Rogoff (1995), Taylor (1995) as well as Breuer (1994) for an overview of the empirical literature. MacDonald (1993) as well as Kugler and Lenz (1993) likewise come to no unequivocal results.

2 See, for example, Frankel and Rose (1995).

3 See Cheung and Lai (1993a) for a very long-term perspective, Masih and Masih (1995) and Soofi (1998) for an analysis of "exotic" currencies as well as Cheung (1993) for an analysis of nominal exchange rates.

4 See, for example, MacDonald (1996), Bayoumi and MacDonald (1999), O'Connell (1998) as well as Taylor and Sarno (1998) for more critical contributions.

they contain such a solution only as a special case. While the empirical results of structural exchange rate models (for example, the monetary model) are rather disappointing, other approaches which incorporate additional fundamentals prove to be more successful.[5] They encompass models with a concrete normative concept of the equilibrium exchange rate; the most notable are the “Fundamental Equilibrium Exchange Rates” (FEER) advanced by Williamson (1994). In this approach the real equilibrium exchange rates are derived in macroeconomic models on the assumption that the economies are simultaneously in their medium to long-term internal and external equilibria. Another strand of literature seeks to explain the actual trend in the exchange rate by estimating reduced-form exchange rate equations. An example of this type is the “Natural Real Exchange Rate” (NATREX) model elaborated by Stein (1994, 1999), in which the regressions are derived from a macroeconomic equilibrium model. Sustainable or desired current account positions no longer play a role in this approach, however, as it is implicitly assumed that the market participants are always in their desired position (Clark and MacDonald 1999). By contrast, the approach used below is based on a still more pragmatic premise in that it does not assume that the determinants are continuously in equilibrium. It builds on earlier empirical studies conducted by the Deutsche Bundesbank on the external value of the D-Mark.[6]

As the currencies of 11 European economies were merged into the euro when it was launched at the beginning of 1999, the following analysis deals with the factors determining the (real) exchange rate of the euro against the US dollar.[7] The analysis is structured as follows: the next section briefly describes the calculation method used for the “synthetic” euro prior to 1999. The core of the analysis is section 3-5, which examines theoretically (section 3) and empirically (section 4) the relationship between a synthetic euro exchange rate and its driving forces and determines the empirical model which is best suited to analyse the exchange rate (section 5). Section 6 contains an assessment of the euro exchange rate since the beginning of the year; the last section summarises the most important results and provides a brief outlook.

2. A “synthetic” euro exchange rate

In view of the fairly short period for which data on the exchange rate trend of the euro are available, an approach geared to the medium term naturally requires some kind of artefact: a substitute for the (real) exchange rate of the euro against the US dollar must be calculated

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5 See the pioneering work in Meese and Rogoff (1983) as well as the overview of the literature in Frankel and Rose (1995) and in MacDonald (1995).

6 See Deutsche Bundesbank (1995) as well as Clostermann and Friedmann (1998). For studies undertaken outside the Bundesbank see also, for example, MacDonald (1999) as well as Habermeier and Mesquita (1999) with a panel approach. As in the NATREX concept, the exchange rates forecast with this approach are sometimes also referred to as the "Behavioral Equilibrium Exchange Rates" (BEER).

7 The terms *effective exchange rate* and *external value* are used synonymously below.

for the past. Such a "synthetic" real euro exchange rate has been calculated as a geometrically weighted average of the dollar exchange rates of individual EMU currencies.[8] The applied weights are those which are provided by the BIS and which represent the share of the foreign trade of the individual euro area countries with non-euro area countries in total EMU trade with non-euro area countries. Until recently the effective exchange rates of the euro published by the ECB were also based on this weighting scheme.[9] As expected, Germany, as the largest economy in the euro area, is given the highest weighting, at 37.4 %, followed by France and Italy, whose weightings, at 18.4 % and 14.2 %, respectively, are less than half as large.

For conceptual reasons this procedure is given preference to the alternative of a weighting pattern which represents the breakdown of bilateral trade between the United States and the euro area and which could actually provide a better means of obtaining an indicator reflecting bilateral competitiveness. However, if the exchange rates of the synthetic euro were always determined by the bilateral trade breakdown between the euro area and the individual non-euro area partner countries, the key transitivity condition for exchange rates would be violated. In other words, by applying bilateral weights the actual yen-dollar exchange rate could not be derived on the basis of the triangular arbitrage condition using the synthetic exchange rate of the euro against the dollar and the synthetic exchange rate of the euro against the yen.

In more formal terms, an index for the synthetic nominal effective euro rate $W_{\$/Euro}$ is then obtained by using the formula:

$$(1) \qquad W_{\$/Euro} = \prod_{i=1}^{10} W_{\$/i}^{\;\gamma_i} \quad .$$

where $W_{\$/i}$ is an index of the effective exchange rate of the EMU currency i against the US dollar and γ_i denotes the corresponding foreign trade weight.[10] By analogy, after elimination of the different price trends at home and abroad the synthetic real effective

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8 Although the ECU has occasionally been regarded as the forerunner of the euro, the use of the ECU/dollar rate does not seem to be appropriate in this context. Firstly, the ECU basket contained currencies which were not merged into the euro at the beginning of the year (the pound sterling, the drachma and the Danish krone) and, secondly, it did not include the currencies of some other countries which are participating in EMU (the Austrian schilling and the Finnish markka). This gives rise to justified doubts about the parallelism of the conceivable euro and ECU developments for the past.

9 The ECB has meanwhile published its own calculations of the effective exchange rate of the euro in the October 1999 issue of its Monthly Bulletin. However, its exchange rate pattern differs only very little from the one used here.

10 The Belgian and Luxembourg francs, which were already components of a monetary union, are regarded as being one currency here.

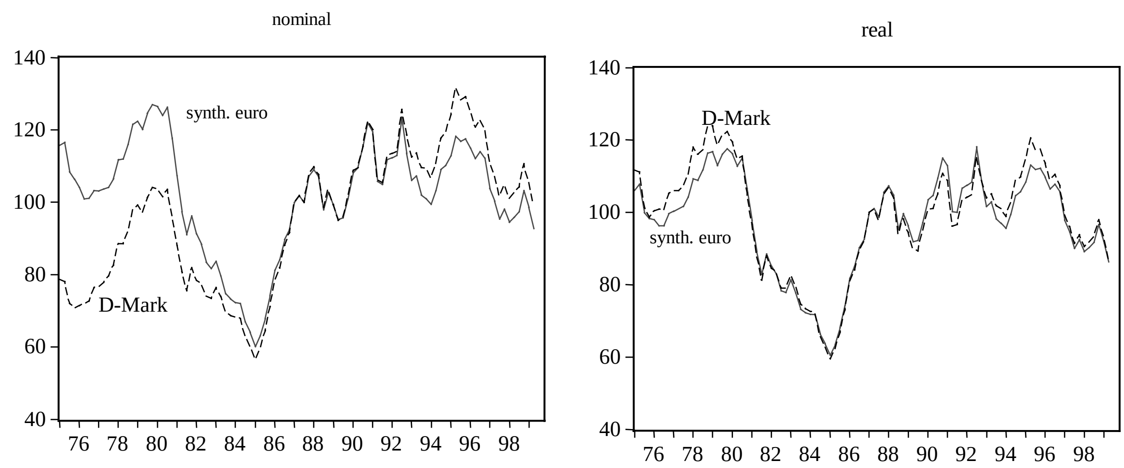
euro rate $Q_{\$/Euro}$ is obtained as a geometric mean of the price-adjusted nominal effective exchange rates of the individual EMU currencies against the US dollar:

$$(2) \qquad Q_{\$/Euro} = P_{US}^{-1}\prod_{i=1}^{10}(W_{\$/i}P_i)^{\gamma_i} \quad .$$

In chart 1 the synthetic effective euro rates against the US dollar are briefly compared in nominal and real terms with the corresponding D-Mark rates. Each rise in the curve represents an appreciation, and the real exchange rates are based on consumer prices.

Chart 1

External values of the D-Mark and
and of the synthetic euro against the US dollar



While the nominal external value of the D-Mark had been showing a tendency to appreciate against the US dollar since the mid-seventies, the synthetic nominal euro rate against the US currency was recently being quoted again at a level that was actually somewhat below that of the second half of the eighties. This development primarily reflects the greater internal price stability of the D-Mark both compared to the US dollar and to a weighted average of the EMU currencies. By contrast, the real external values of the euro and the D-Mark move very similarly to one another (correlation coefficient > 98 %), which indicates to a certain extent the tendency to balance out the inflation differential in Europe by nominal exchange rate adjustments.

3. Determinants of the "synthetic" real euro exchange rate

3.1. Theoretical concept

Theoretically, the uncovered interest rate parity should be a major determinant of a currency's exchange rate. Ignoring risk premia and assuming free cross-border capital movements, the arbitrage equilibrium on the international foreign exchange markets requires that interest rate differentials (i*-i) between foreign and domestic markets equal the exchange rate changes of the home currency ($\Delta w = \ln(\Delta W)$) expected at time t for the maturity t+k. Denoting variables relating to foreign markets with an asterisk and using E as the expectation operator yields:

(3) $$E_t(\Delta w_{t+k}) = i_t^* - i_t \ .$$

Subtracting in (4) the expected inflation differential $E_t(\Delta p_{t+k} - \Delta p^*_{t+k})$ between the domestic market and markets abroad on both sides of (3) gives the real interest rate differential in (5):

(4) $$E_t(\Delta w_{t+k} + \Delta p_{t+k} - \Delta p^*_{t+k}) = i_t^* - i_t + E_t(\Delta p_{t+k} - \Delta p^*_{t+k}) \ .$$

(5) $$E_t(\Delta q_{t+k}) = r_t^* - r_t \ .$$

Furthermore, agents expect that in the long term the real exchange rate ($q = \ln(Q)$) will revert to its equilibrium value $\overline{q}$ only gradually. This gives the following equation (6), in which the (expected) speed of adjustment to equilibrium is determined by θ:

(6) $$E_t(\Delta q_{t+k}) = -\theta(q_t - \overline{q}_t) \ , \quad \text{where } 0 < \theta < 1 \ .$$

Combining (5) and (6) yields:

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

which implies that the real interest rate differential ensures that over the medium term the current real exchange rate reverts to its equilibrium level.

As mentioned above, PPP entails a constant real long-term equilibrium exchange rate $\overline{q}$. However, this assumption may be too strong. Balassa (1964) and Samuelson (1964), in particular, have shown that different sectoral productivity trends at home and abroad can have an impact on the real equilibrium exchange rate based on broad price and cost indices. According to their model, relatively larger increases in productivity in the traded goods sector are accompanied *ceteris paribus* by a systematic tendency of a country's currency to appreciate. This notion will be outlined more formally in the following.[11]

In logarithmic terms, the price level in a country is derived from the weighted average of the prices of traded goods (T) and non-traded goods (N), where α is the share of traded goods in the domestic basket of goods and β the share of traded goods in the foreign basket of goods:

(8a) $$p_t = \alpha p_t^T + (1-\alpha) p_t^N ,$$

(8b) $$p_t^* = \beta p_t^{T*} + (1-\beta) p_t^{N*} .$$

In addition, the real exchange rate is defined as

(9) $$q_t \equiv w_t + p_t - p_t^* \quad .$$

Allowing for the arbitrage condition for internationally traded goods,

(10) $$q_t^T \equiv w_t + p_t^T - p_t^{T*} \quad ,$$

yields equation (11) by inserting (8a), (8b) and (10) into (9) and rearranging:

(11) $$\bar{q}_t = q_t^T + (1-\alpha) \cdot (p_t^N - p_t^T) - (1-\beta) \cdot (p_t^{N*} - p_t^{T*}) \; .$$

Assuming productivity-based wages and a fixed mark-up (φ), the relative prices within a given country move inversely to the ratio of productivity gains in these two sectors. Based on Cobb-Douglas technology, the following equation is obtained at given international interest rates:[12]

(12) $$(p_t^N - p_t^T) = \varphi((\eta^N / \eta^T) h_t^T - h_t^N) ,$$

where h is the total factor productivity in each sector and η the output elasticity of the labour input in the production function. An increase in productivity in the traded goods sector therefore results *ceteris paribus* in a proportional fall in relative prices between this sector and the non-traded goods sector implying a real appreciation.

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[11] See Deutsche Bundesbank (1995), Clostermann and Friedmann (1998) as well as HypoVereinsbank (2000) for a similar account of the interrelationships.

[12] See Asea and Corden (1994) for a detailed account as well as Chinn (1999) for a similar presentation.

Portfolio equilibrium models are another possible channel, through which fundamental factors may affect the real exchange rate. They presume that an increase in external debt, which is equal to the current account deficit, is absorbed by internationally diversifying investors only if the adjustment in their portfolio structures is rewarded by a higher yield. At a given interest rate, this is accomplished through a depreciation of the currency of the debtor country, which in turn strengthens the country's international competitiveness, boosting exports, which in turn can be used to service interest payments. In the present study, however, limitations regarding the availability of data prevented the inclusion of the net external asset position between the euro area and the United States. Substituting this variable by an accumulative current account for the euro area which the OECD provides on an annual basis from 1975 onwards did not lead to any significant results.[13] This may, however, also be due to data and valuation problems.

Assuming that higher government spending may undermine long-term confidence in a currency, a reaction of the real exchange rate to a change in government expenditure could also be expected. This argument is based on the supposition that agents suspect that higher spending today could be followed by distorting taxes or monetisation of government debt in the future, with both adversely affecting economic activity.[14] Barro and Lee (1994) examined empirically the relationship between government spending and growth in a comprehensive cross-sectional analysis and found that increased government expenditure is accompanied by lower aggregate productivity growth.[15] In line with MacDonald (1999) the model is therefore extended by adding the relative expenditure ratio between the euro area and abroad (g_t - g_t*), which, at least in the long term, should have a negative impact on the real exchange rate.[16]

$$(13) \qquad \overline{q}_t = q_t^T + (1-\alpha)(p_t^N - p_t^T) - (1-\beta)(p_t^{N*} - p_t^{T*}) - \phi(g_t - g_t^*) \quad .$$

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13 See OECD, A comprehensive data set for the euro area, which is available on the Internet at http://www.oecd.org/eco/data/euroset.html.

14 See Habermeier and Mesquita (1999) for this argument. If a country's current account position is a reflection of its fiscal behaviour, for example, when an expansionary fiscal policy fosters the country's borrowing abroad, a variable that reflects government action ought to show a notable degree of similarity with the net external assets position.

15 Strictly speaking, however, government expenditure in their analysis is adjusted to eliminate, for example, educational expenditure as this variable might have a direct impact on productivity. Such a breakdown was not undertaken here.

16 On the assumption that government demand for non-traded goods would tend to be relatively stronger than private demand for this category, one might suppose in the short term that an expansionary government spending policy would lead to rising prices in this sector and therefore to a real appreciation of the currency. See, for example, Chinn (1997), who implemented this argument formally in a real exchange rate model.

Furthermore, exogenous terms of trade shocks such as a permanent change in oil prices may also be reflected in the relative prices of traded goods q_t^T.[17] In general, permanent increases in oil prices improve the international competitiveness of those countries which are relatively less dependent on oil imports (or which actually export oil). C*eteris paribus* an oil price increase should therefore result in a real appreciation of the currency of the country less dependent on oil. Although the United States consumes more oil relative to its economic activity than the euro area countries, the United States is at the same time more self-sufficient in oil whereas the euro area countries are almost entirely dependent on imports of oil to meet their needs. Consequently, a permanent rise in real oil prices should result in a real depreciation of the synthetic euro.

In summary, then, the following equation for the real exchange rate can be obtained by combining (11) and (7):

$$(14) \qquad q_t = q_t^T[p^{oel}] + (1-\alpha)(p_t^N - p_t^T) - (1-\beta)(p_t^{N*} - p_t^{T*}) - \phi(g_t - g_t^*) - (1/\theta)(r_t^* - r_t)\,.$$

3.2. Data

Quarterly data covering the period from the first quarter of 1975 to the fourth quarter of 1998 are used for the econometric analysis. This period corresponds to the era of floating exchange rates between the euro area countries and the United States after a brief "adjustment phase" following the collapse of the Bretton Woods system. A longer estimation period covering the Bretton-Woods period of fixed exchange rates as well seems to be rather problematic since it is difficult to believe that the underlying data-generating processes are the same in both regimes (Chinn 1997). As usual, all variables with the exception of the real interest rate differential are included in the estimates in logarithms. The real exchange rate (q) of the euro against the US dollar has been measured as in formula (2). Consumer prices are used to adjust the nominal exchange rate for different inflation histories at home and abroad because of the high data quality and because this price index is currently being used by the ECB to calculate the real exchange rate of the euro. Nevertheless, a broader price index such as the GDP deflator or the price index of total sales might be a better representation of the basket of goods comprising traded and non-traded goods. As soon as such data are available, the calculations are also to be made on the basis of these deflators. However, analyses for the D-Mark have indicated that the deflator of total sales and consumer prices show fairly similar trends (Clostermann and Friedmann 1998), and therefore the makeshift use of the consumer price indices should not have too detrimental an effect on the results.

[17] Changes in the terms of trade can also affect the real exchange rate through wealth effects, shifts in intertemporal consumer preferences or a change in the production function (Chinn and Johnston 1997).

To take account of Balassa's productivity bias of real exchange rates, the relative prices of traded and non-traded goods in the euro area compared with the same ratio for the United States (tnt) are chosen. Producer price indices are used as proxies for the prices of traded goods and consumer price indices as proxies for the prices of non-traded goods.[18] In theory, a rise in this variable should result in a real appreciation of the synthetic euro. Alternatively, evidence of a more direct productivity variable – approximated, for instance, by the ratio of GDP to the number of employed persons – has also been examined. Although this variable was found to be important in the estimates by Clostermann and Friedmann (1998) for the D-Mark, it has been consistently insignificant in the present estimates for the euro area. Faruqee (1995) and Feyzioglu (1997) have tried to model the productivity trend directly through sectoral data on the total factor productivity. However, these data are not available on a quarterly basis and are associated with problems of homogeneity across countries (see Alberola et al. (1999)).

The weighted spot market dollar price for petroleum is used, after adjustment for the trend in producer prices in the United States, as the real oil price (oel). When used as an alternative, dollar-based HWWA index of raw material prices has always given less satisfactory empirical results. The real interest rate differential (rze) between the euro area and the United States is determined from the difference between a weighted European long-term government bond yields and the US equivalent (IFS statisitcs), adjusted for annual consumer price inflation. By contrast, short-term real interest rates have regularly produced less satisfactory test results. The ratio of government expenditure to GDP in the euro area relative to the ratio of government expenditure in the United States is used as a fiscal variable (fisc). All aggregates for the euro area are constructed in line with the weighting procedure for calculating the synthetic real euro rate.[19]

Specifically, the following approach is therefore estimated for the real exchange rate (the expected signs are given in brackets):

$$q_t = f(\overset{(+)}{rze}, \overset{(+)}{tnt}, \overset{(-)}{oel}, \overset{(-)}{fisc}) \ .$$

---

[18] Although both the producer prices and the consumer prices contain prices for traded and non-traded goods, the proportion of non-traded goods in the basket of goods used to calculate producer prices is considerably smaller than in the basket of goods used to calculate consumer prices. See also Engel (1999) for a critical discussion of various means of measuring the impact of non-traded goods on the real exchange rate.

[19] In line with Chinn and Johnston (1997) the empirical part included an examination to ascertain if per capita income in the euro relative to that of the United States could explain shifts in consumer preferences. However, this variable did not improve the results.

4. Econometric analysis

4.1. ADF tests

The ADF tests suggest (see appendix) that the majority of the variables are integrated in levels and stationary in first differences. The only exception is the real interest rate differential which (in line with theoretical considerations) is identified as a stationary variable.[20] There are some indications of heteroscedastic residuals in the underlying regression for the relative price relationship between traded and non-traded goods (tnt). However, a non-parametric Phillips-Peron test which was therefore conducted produces results which are largely identical with those of the ADF test. Consequently, it is assumed that the variable is I(1). Apart from that, LM and ARCH tests do not identify any indications of auto-correlation or heteroscedasticity in the residuals.

4.2. Johansen procedure

Owing to the non-stationarity of the time series, the euro-dollar exchange rate is estimated in a vector error correction model (VECM) based on the procedure developed by Johansen (1991) as well as Johansen and Juselius (1994). This approach seems to be particularly suited to verify the long-term equilibrium relationships (= cointegration relationships) on which the theoretical considerations are based. The empirical analysis starts with an unrestricted vector error correction model which takes the following form:

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

where y_t represents the vector of the non-stationary variables q_t, tnt_t, oel_t, and $fisc_t$. ε_t denotes the vector of the iid residuals and μ the vector of the constant terms. The number of cointegration relationships corresponds to the rank of the matrix Π. If the matrix Π has a reduced rank ($0 < r < p$), it can be separated into a (p x r)-dimensional matrix of the loading coefficients α and a (p x r)-dimensional matrix of the cointegration vectors β ($\Pi = \alpha\beta'$). The cointegration vectors represent the long-term equilibrium relationships in the system. The loading coefficients denote the importance of the cointegration relationships in the individual equations and the speed of adjustment following deviations from the long-term equilibrium. In the first step the lag order (=k) of the system is determined by estimating an unrestricted VAR-model in levels and using the information criteria proposed by Akaike and, alternatively, by Hannan-Quinn, which is usually more restrictive

[20] Edison and Melick (1999) also find that the real exchange rate should be treated as a I(1) variable and the real interest rate differential as a I(0) variable. MacDonald (1998) as well as Clostermann and Friedmann (1998), by contrast, find empirical support to the real interest rate differential as a I(1) variable.

with regard to the lag structure to be chosen. Both criteria recommend a lag length of 2, for which the residuals are not auto-correlated.

The number of cointegration vectors is verified in a second step by determining the cointegration rank with Reimers' adjusted trace test.[21] This test suggests precisely one cointegration relationship, i. e. one equilibrium relationship between the non-stationary variables q, tnt, oel, and fisc. (see table 1).

Table 1
Test for the number of cointegration relationships in the VECM

Null hypothesis	Trace test (with Reimers' adjustment)	Critical values (95%)
r<1	53.92	47.21
r<2	27.27	29.68
r<3	8.40	15.41
r<4	0.20	3.76

It therefore seems reasonable to restrict the VECM to one cointegration relationship and – as the real interest rate parity suggests – to include the real interest rate differential as a stationary (exogenous) variable (with a lag length of 0 to 2). Hence, a VECM with the following structure is estimated:

$$\begin{pmatrix} \Delta q_t \\ \Delta tnt_t \\ \Delta oel_t \\ \Delta fisc_t \end{pmatrix} = \Gamma_1 \begin{pmatrix} \Delta q_{t-1} \\ \Delta tnt_{t-1} \\ \Delta oel_{t-1} \\ \Delta fisc_{t-1} \end{pmatrix} + \begin{pmatrix} \alpha^{q} \\ \alpha^{tnt} \\ \alpha^{oel} \\ \alpha^{fisc} \end{pmatrix} \begin{pmatrix} 1 & \beta^{tnt} & \beta^{oel} & \beta^{fisc} \end{pmatrix} \begin{pmatrix} q_{t-1} \\ tnt_{t-1} \\ oel_{t-1} \\ fisc_{t-1} \end{pmatrix} + \psi \begin{pmatrix} rze_t \\ rze_{t-1} \\ rze_{t-2} \end{pmatrix} + \mu + \varepsilon_t \quad .$$

The long-term relationship of this system – after the cointegration coefficients have been normalised to the real exchange rate – is obtained from

21 Reimers (1992) points out that the critical values identified by means of trace statistics can depend on the data-generating process and the sample length. He proposes an adjustment of the calculation for the number of endogenous variables (p) and the lag length (k). This results in the following, modified trace statistics:

"Conventional" trace test: $LR = -T \sum_{i=r+1}^{p} \log(1-\lambda_i)$.

Reimers' adjusted trace test: $LR = -(T - pk) \sum_{i=r+1}^{p} \log(1-\lambda_i)$.

$$q - \beta^{tnt} \cdot tnt - \beta^{oel} \cdot oel - \beta^{fisc} \cdot fisc \, ,$$

where the β values reflect the long-term coefficients.

To interpret the long-term relationship as an exchange rate equation, however, all variables except the exchange rate must meet the condition of weak exogeneity, i.e. deviations from the long-term equilibrium are corrected solely through exchange rate responses. As mentioned above, the extent to which the individual variables adjust to the long-term equilibrium is expressed in the α-values. In a formal test, the hypothesis of weak exogeneity of tnt, oel and fisc ($\alpha^{tnt}=\alpha^{oel}=\alpha^{fisc}=0$) cannot be rejected at the standard levels ($\chi^2(3)= 6.36$). Taking this restriction into account, the following regressions for the VECM ensue:

Table 2
Regressions and test statistics of the VECM
(t-values in parentheses)

$$\Delta q_t = -\underset{(4.38)}{0.27} \cdot (q_{t-1} - \underset{(3.12)}{2.78}\, tnt_{t-1} + \underset{(4.06)}{0.34} \cdot oel_{t-1} + \underset{(3.29)}{0.94} \cdot fisc_{t-1})$$

$$+ \underset{(2.87)}{0.28} \cdot \Delta q_{t-1} - \underset{(0.80)}{0.47} \cdot \Delta tnt_{t-1} + \underset{(2.00)}{0.08} \cdot \Delta oel_{t-1} + \underset{(2.15)}{0.62} \cdot \Delta fisc_{t-1}$$

$$+ \underset{(2.80)}{1.52} \cdot rze_t + \underset{(0.33)}{0.25} \cdot rze_{t-1} - \underset{(1.01)}{0.61} \cdot rze_{t-2} + \underset{(4.38)}{1.62}$$

$$\Delta tnt_t = +\underset{(1.57)}{0.03} \cdot \Delta q_{t-1} - \underset{(1.89)}{0.23} \cdot \Delta tnt_{t-1} + \underset{(3.31)}{0.03} \cdot \Delta oel_{t-1} + \underset{(2.86)}{0.17} \cdot \Delta fisc_{t-1}$$

$$+ \underset{(0.63)}{0.07} \cdot rze_t + \underset{(0.13)}{0.02} \cdot rze_{t-1} - \underset{(0.19)}{0.02} \cdot rze_{t-2} - \underset{(0.01)}{0.00}$$

$$\Delta oel_t = -\underset{(0.07)}{0.02} \cdot \Delta q_{t-1} - \underset{(3.26)}{5.79} \cdot \Delta tnt_{t-1} + \underset{(2.86)}{0.36} \cdot \Delta oel_{t-1} + \underset{(0.36)}{0.31} \cdot \Delta fisc_{t-1}$$

$$+ \underset{(2.16)}{3.52} \cdot rze_t - \underset{(2.57)}{5.75} \cdot rze_{t-1} + \underset{(1.97)}{3.54} \cdot rze_{t-2} - \underset{(0.01)}{0.01}$$

$$\Delta fisc_t = -\underset{(1.02)}{0.04} \cdot \Delta q_{t-1} + \underset{(1.60)}{0.36} \cdot \Delta tnt_{t-1} - \underset{(1.41)}{0.02} \cdot \Delta oel_{t-1} - \underset{(0.32)}{0.03} \cdot \Delta fisc_{t-1}$$

$$- \underset{(0.42)}{0.09} \cdot rze_t + \underset{(0.31)}{0.09} \cdot rze_{t-1} + \underset{(0.33)}{0.08} \cdot rze_{t-2} + \underset{(0.02)}{0.00}$$

Test statistics	Equation: Δq	Equation: Δtnt	Equation Δoel	Equation Δfisc
R^2 (adj.)	0.31	0.17	0.14	-0.01
Standard error (in %)	3.78	0.75	1.12	1.40
LM(1)	0.26	0.05	1.25	3.54 *
LM(4)	0.13	1.16	1.29	1.25
ARCH(1)	0.89	0.01	0.00	0.48
ARCH(4)	0.37	0.17	1.47	0.80
Jarque-Bera	2.47	3.89	63.93 ***	0.92
***(**,*): Error probability < 1% (5%, 10%).				

The term in parentheses in the first equation (=Δq) describes the long-term relationships. The variables show the expected signs, and the coefficients are highly significant according to their standard errors. In the long run, an increase in the relative price ratio, a decrease in real oil prices and in relative government expenditure result in a real appreciation of the euro exchange rate.

4.3. Single equation approaches

Owing to the weak exogeneity of the three fundamentals, switching to a single equation error correction model (=SEECM; Engle et al. 1983, Johansen 1992), may still improve the efficiency of the estimate. In line with Banerjee et al. (1998) the following equation is specified:

$$\Delta q_t = \alpha \cdot (q_{t-1} - \beta^{tnt} \cdot tnt_{t-1} - \beta^{oel} \cdot oel_{t-1} - \beta^{fisc} \cdot fisc_{t-1}) + \sum_{i=1}^{k} (\delta_i^{q} \cdot \Delta q_{t-i}) + \sum_{i=-m}^{k} (\delta_i^{tnt} \cdot \Delta tnt_{t-i} + \delta_i^{oel} \cdot \Delta oel_{t-i} + \delta_i^{fisc} \cdot \Delta fisc_{t-i} + \delta_i^{rze} \cdot rze_{t-i})$$

The first part of the expression with the variables in levels describes the cointegration relationship that has been normalised to the exchange rate. The SEECM differs from the VECM in that contemporary variables as well as leads of the (weakly) exogenous variables are added in order to improve the asymptotical properties of the estimate.[22] The lead length (=m) is restricted to a maximum of two as recommended by Banerjee et al. (1998, p. 275); the lag length (=k) is restricted to a maximum of four. The regression has been run with the so-called backward procedure, i. e. insignificant coefficients (error probability > 5 %) have been successively omitted (Hendry and Richard 1983). The final regression is thus:

[22] These approaches are based on Phillips and Lorethan (1991) as well as Saikkonen (1991). Chinn and Johnston (1997) also provide an example of how the approaches are applied to empirical exchange rate analysis.

Table 3
Regression and test statistics of the SEECM
(t-values in parentheses)

$$\Delta q_t = -\underset{(5.10)}{0.28} \cdot (q_{t-1} - \underset{(1.94)}{1.92}\, tnt_{t-1} + \underset{(3.04)}{0.24} \cdot oel_{t-1} + \underset{(2.23)}{0.64} \cdot fisc_{t-1})$$

$$+ \underset{(2.75)}{0.24} \cdot \Delta q_{t-1} + \underset{(4.10)}{2.02} \cdot \Delta tnt_t + \underset{(3.98)}{1.30} \cdot rze_t + \underset{(4.66)}{1.58}$$

Test statistics	
R^2 (adj.)	0.40
Standard error (in %)	3.61
LM(1)	0.08
LM(4)	0.12
ARCH(1)	1.52
ARCH(4)	0.88
Jarque-Bera	0.57
Cusum	stable
Cusum Square	stable
***(**,*): error prob. < 1% (5%, 10%).	

The coefficients of the long-term relationship show the anticipated signs and are statistically significant at standard levels. The dimension of the coefficients largely resemble those of the long-term equation in the Johansen procedure. This suggests that the relationship implies some stability irrespective of the applied econometric methodology. Thus, a lasting rise in the price ratio between traded and non-traded goods in the euro area in relation to the United States has a positive effect on the real synthetic euro exchange rate. Conversely, a permanent rise in (real) oil prices results in a weakening of the real exchange rate of the euro, as the euro area countries are relatively more dependent on raw materials. The coefficient of 0.24 indicates that a permanent rise in real oil prices of, say, 10 % will result in a lasting depreciation of the euro exchange rate of 2.4 %. The (contemporarily estimated) semi-elasticity of the real interest rate differential is positive and highly significant. In the short run, an increase of 1 percentage point in the real interest rate differential results in a 1.3 % increase in the real effective euro exchange rate. This value, being greater than 1, implies that the real exchange rate is on average more volatile than the real interest rate differential. This may be a consequence of price rigidities, which can be reflected in overshooting exchange rates in the short term. The significantly positive relationship between the real exchange rate and its first lag could be interpreted as an indication that the exchange rate is also being driven, in the short run, by non-fundamental factors. This could be due to the market behaviour of chartists, whose exchange rate forecasts are customarily based to some extent on past exchange rate movements (Nagayasu 1999).

The coefficient of the error correction term is negative and, measured against the critical values provided by Banerjee et al. (1998), highly significant. Thus the condition for a long-term stable equilibrium is satisfied. In comparison with the simple PPP approaches, the inclusion of these additional variables improves the adjustment properties of the model considerably. The parameter value of 0.28 suggests a half life period of shocks of just under two quarters. In other words, the differential between the real exchange rate and the (more flexible) equilibrium is reduced by half in less than two quarters after an exogenous shock.[23] This result is especially remarkable against the background of half life periods of four to five years established in the above-mentioned empirical studies on the PPP.[24]

Breusch-Godfrey Lagrange multiplier tests (LM) give no indication of autocorrelation in the residuals (1st and 4th order). Nor can the Lagrange multiplier test for autoregressive conditional heteroscedasticity (1st and 4th order) identify any violation of the basic assumptions of the regression. In addition, the Jarque-Bera test confirms the normal distribution assumption for the residuals, and CUSUM tests give no indication of parameter or variance instability.

5. Predictive ability

In order to underscore the relative superiority of the single equation error correction model (=SEECM), its forecasting ability is compared with four alternative approaches:

1. **Random Walk model (=RW).** Following the influential article of Meese and Rogoff (1983), this model has become a very popular benchmark for the quality of exchange rate models. In line with the stationarity tests the model is specified without a constant and a trend.[25]
2. **MA(1) model (=MA).** In addition, an ARMA model is included because it is an example of a simple univariate forecast model. The MA(1) structure results from a search process where, on the basis of an ARMA(4,4) model, the most parsimonious specification at which the SIC reached its (absolute) maximum in a regression which covered the complete estimation period is chosen.
3. **VAR model** (k=1) with the variables in first differences **(=DVAR).** This approach enables an assessment to be made as to whether taking account of the cointegration restrictions results in greater efficiency.
4. **VEC model** (k=2) analogous to the approach specified in section 4.2 **(=VECM).** This alternative analysis ascertains whether the loss of information resulting from reducing

[23] The half life period is calculated as $\log(0.5)/\log(1-\alpha)$.
[24] See the comprehensive overview provided by Froot and Rogoff (1995) and Obstfeld (1998).
[25] A Random Walk specification in the form of a "no change" forecast provides nearly identical results.

the four-equation system to a single equation approach is outweighed by the greater efficiency resulting from the smaller number of parameters to be estimated.

For each of these models the real exchange rate is estimated over a period from the first quarter of 1975 to a time t and dynamically forecast for up to h quarters into the future ($=q_{t+h,t}$). All models are based on completely endogenous structures; thus the forecast includes only information that had actually been available at the time it was carried out. In the case of both the VEC and the VAR model the exchange rate forecasts therefore always include forecast values for the other variables which have been generated inside the model itself, while the coefficients of the original estimation periods are kept constant. In both models forecasts of the real exchange rate are generated on the assumption of a constant real interest differential. In the same manner the exogenous variables of the SEECM are replaced, for forecasting purposes, by the last available observation in the respective estimation period.[26]

The forecast error of a model ($=e_{t+h,t}$) is then calculated as the difference between the actual value at time t+h ($=q_{t+h}$) and its forecast value ($q_{t+h,t}$):

$$e_{t+h,t} = q_{t+h} - q_{t+h,t} \quad .$$

The estimates are carried out recursively, the “first” estimation period being from the first quarter of 1975 to the fourth quarter of 1987. On the one hand, this allows sufficient degrees of freedom in the regressions from the outset; on the other hand, it yields a number of forecast values high enough to be corroborated statistically. The specification is then successively extended quarter by quarter, and for each estimation period forecasts ranging from one to eight quarters are generated.

The quality of the forecast values of the competing models can be assessed using measures of forecast accuracy. Therefore, the frequently used root mean squared error (RMSE) is applied as a summary measure:[27]

$$RMSE_h = \sqrt{\frac{1}{T}\sum_{t=1}^{T} e_{t+h,t}^2} \quad .$$

26 Alternative short-term forecasts of the exogenous variables using ARMA processes did not produce any better results.

27 Test calculations on the basis of the mean absolute forecast error (MAE) produced identical results in terms of quality.

A smaller RMSE indicates that a given model has a better predictive ability. The forecast quality of each approach is then measured relative to the RMSE of the SEECM. Consequently, ratios smaller than 1 indicate a lower RMSE of the SEECM compared with the respective competing model and thus a superior predictive quality. Over and above that, a formal test based on the so-called loss differential (Diebold 1998) may provide some insight into whether a model can generate significantly better forecasts from a statistical point of view. Assuming a quadratic loss function, the loss differential is defined as the squared forecast error of the benchmark model (the SEECM in our case) minus the squared forecast error of the alternative model. The test is pursued by regressing the loss differential solely on a constant term. If the constant term is significantly lower than zero it could be concluded that the benchmark model provides *significantly* better forecasts than the other.

The results of these calculations are summarised in table 4. The first column shows the ratio of the two RMSEs. The significance level is indicated by asterisks in the second column. Since the underlying regressions often contain auto-correlated residuals, MA(1) terms have been added in these cases. This has eliminated first and forth order auto-correlation in all instances. Alternatively, regressions are run by adjusting the test statistics with the Newey-West procedure, producing consistent estimates even in the case of auto-correlation and heteroscedasticity (column 4). However, the results are virtually the same regardless of whether the auto-correlated residuals have been adjusted by MA processes or whether the Newey/West (NW) correction has been applied.

Table 4
Forecast quality of different models

Forecast horizon, Quarters	SEEC ↔ RW			SEEC ↔ MA(1)			SEEC ↔ DVAR			SEEC ↔ VECM		
	Ratio	Sig.	NW	Ratio	Sig.	NW	Ratio	Sig.	NW	Ratio	Sig.	NW
1	1.01			1.01			0.90			0.92		
2	0.95	1)		0.92			0.84		*	0.87	*	
3	0.83	1)		0.81	1)		0.77	**	**	0.85	1)	
4	0.79	**1)	*	0.79	**1)	*	0.74	**	**	0.82	*1)	*
5	0.75	***1)	**	0.73	***1)	***	0.68	***	***	0.80	**1)	**
6	0.72	***1)	***	0.69	***1)	***	0.65	***	***	0.77	**1)	**
7	0.73	***1)	**	0.70	***1)	***	0.66	***	***	0.76	**1)	**
8	0.74	**1)	***	0.72	***1)	***	0.67	***	***	0.73	***	**

***/**/* Error probability 1%/5%/10%. 1) MA(1) process

With respect to exchange rate forecasts for the next quarter, the SEECM cannot outperform the Random Walk-Model and the MA(1) process. This can be seen from RMSE ratios, which are greater than 1 for the first quarter. Nevertheless, the test statistics do not suggest that the forecasting performance of these two alternative models are also in statistical terms significantly better than the SEECM.

Beyond a horizon of one quarter the predictive ability of the SEECM improves significantly relative to all other models. For all competing models the sum of the squared forecast errors increases rather quickly for longer horizons, while the forecasting performance of the SEECM deteriorates only very slowly. In the table, this is reflected in rapidly diminishing ratios. Thus the SEECM outperforms all competing models even if the forecasts cover as little as two quarters. The comparatively higher predictive ability of the SEECM compared with the DVAR is therefore another indication of the validity of the established cointegration relationship. Moreover, the VEC model does not achieve lower RMSEs than the SEECM for any of the forecasting horizons examined.

The statistical tests based on the loss differentials underpin the relative superiority of the SEECM for longer-term forecasting horizons. Given the usual levels of significance, the SEECM significantly outperforms all of the competing models even for exchange rate forecasts covering less than one year. For forecasts covering periods of more than one year, the SEECM's superiority is even more evident as the relative RMSEs continue to decrease and its significance increases. Measured against these criteria, the VEC model is significantly outperformed by the SEECM for longer forecasting periods. Hence, the reduction of the VEC model to a single equation approach is an efficient strategy for improving the estimation results in this case.

6. Recent euro developments

Since the beginning of 1999 the euro has weakened considerably against the US dollar. Only a few days after its launch, the new European currency began a downward trend that continued until mid-July. During this time the euro weakened from its peak level of US$ 1.17 to US$ 1.01. The new European currency then recovered somewhat against the dollar and fluctuated within a narrow range of between US$ 1.03 and US$ 1.07. Only recently, however, the value of the euro declined again and was quoted below parity against the US dollar.

The SEECM presented above can now be used to assess these developments. To this end, it was estimated until the end of 1998. The structural links which were thus identified were then used to gain some insight into the exchange rate developments that this model would have predicted using the available data for the first three quarters of 1999. According to these simulations, the current equilibrium exchange rate would be approximately 1.20 $/€. One reason for this prediction of an upward trend since the beginning of 1999 is that the real interest rate differential was reversed in favour of euro-denominated paper in the autumn of last year, a development which had a positive effect on the predicted exchange rate of the euro. A more important factor, however, is the decline in oil prices until the first quarter of 1999. Since oil price changes feed only slowly into real exchange rate adjustments, these effects are still felt in the following quarters. They seem to be strong enough to disguise the counteracting effects of the surging oil prices since February 1999, which will cause the equilibrium exchange rate to depreciate in due course. This can be seen in forecasts beyond the third quarter of 1999, for which the data for the exogenous variables are held constant. The remaining dynamics in the model forces the estimated equilibrium exchange rate of the euro over the medium term to a level of roughly US$ 1.126.

However, such analyses should not be taken too literally. The extent to which such a model is suitable for determining equilibria in the foreign exchange markets and therefore for examining current exchange rate developments can only be assessed by taking into account the margin of uncertainty associated with the empirical approach. The test statistics of the SEECM already indicate, however, that the error margin, at an adjusted R-squared of less than 50 %, is still relatively high. Thus the model can hardly be expected to provide precise short-term exchange rate forecasts. On the 95 % level the confidence interval covers a considerable span, namely, as much as 26 cents in nominal terms, ranging in the medium perspective from below parity to 1.259 $/€. This considerable margin of uncertainty indicates that this method must not be used to assess the exchange rate relationships that are too schematically based on the forecast means.

7. Concluding remarks and outlook

Assuming that exchange rate developments of the euro can be adequately approximated with the aid of a synthetic exchange rate, which has been computed as the weighted average of the currencies making up the euro, the findings could be summarised by a quotation from Obstfeld (1995): “... although short-run fluctuations remain mysterious, the theory is not without predictive content ...” In particular, the following results are worth noting:

1. Looking beyond the horizon of the PPP theory, there are some promising approaches to explain systematically deviations of the exchange rate from the trend mapped out by inflation differentials – at least over the medium term.
2. In addition to the real interest rate differential, it was particularly real oil prices, the relative price relationship between traded and non-traded goods as well as the public sector spending ratio in relation to that of the United States that was identified as

determinants of real exchange rate movements. If these factors are taken into account in the empirical analysis, the pace of adjustment is considerably greater than approaches based on the PPP. The half life of an adjustment to the fundamental-related equilibrium is reduced considerably. It is approximately six months, compared with the "rule of thumb" of four to five years identified in empirical analyses relating to the theory of the PPP.

3. A comparison of the forecasting quality of different approaches shows that the single equation error correction model delivers the best results in the medium term and that it even outperforms the Random Walk model – which has so far been the benchmark for empirical exchange rate models.
4. Over the short term, however, the calculations are still subject to considerable margins of error with the result that great caution should be taken when using this method to assess current exchange rate relationships or daily fluctuations.

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Appendix

Stationarity tests[28]

LEVELS									
Variable	lag AIC	Type	ADF test		LM(1) (p-value)	LM(4) (p-value)	ARCH(1) (p-value)	ARCH(4) (p-value)
q	1	N	-0.218		0.785	0.243	0.243	0.698
tnt	3	N	-1.443		0.999	0.790	0.032	0.210
oel	4	N	-0.726		0.310	0.475	0.760	0.858
fisc	0	C	-2.421		0.695	0.626	0.444	0.186
rze	5	N	-3.043	**	0.999	0.994	0.598	0.938
DIFFERENCES								
Δq	0	N	-7.016	**	0.830	0.237	0.234	0.692
Δtnt	2	N	-4.909	**	0.748	0.641	0.099	0.395
Δoel	3	N	-5.282	**	0.475	0.549	0.680	0.816
Δfisc	0	N	-9.080	**	0.108	0.251	0.429	0.191
**/* = error probability 1 %, 5 %								

28 Augmented Dickey-Fuller tests (ADF tests) are run to determine the order of integration of the variables since integrated variables require a different empirical treatment from stationary variables due to the well-known problem of spurious regression. A sequential testing procedure is chosen by initially establishing the number of lags to be included with the Akaike information criterion, starting with a maximum lag length of 8. Then, three alternative specifications are tested: firstly, with a constant and a trend (T), secondly, with a constant only (C) and, thirdly, without a constant and a trend (N). In each step the significance of the trend and/or the constant and the hypothesis of a unit root is verified in a joint test in order to determine the appropriate specification. Moreover, the presence of auto-correlation (Breusch-Godfrey test) or heteroscedasticity (ARCH test) in the residuals is assessed.

The following papers have so far been published:

May	1995	The Circulation of Deutsche Mark Abroad	Franz Seitz
June	1995	Methodology and technique for determining structural budget deficits	Gerhard Ziebarth
July	1995	The information content of derivatives for monetary policy – Implied volatilities and probabilities	Holger Neuhaus
August	1995	Das Produktionspotential in Ostdeutschland *	Thomas Westermann
January	1996	Sectoral Disaggregation of German M3	Vicky Read
March	1996	Monetary aggregates with special reference to structural changes in the financial markets	Michael Scharnagl
March	1996	The impact of interest rates on private consumption in Germany	Hermann-Josef Hansen
May	1996	Market Reaction to Changes in German Official Interest Rates	Daniel C. Hardy
May	1996	The role of wealth in money demand	Dieter Gerdesmeier

* Available in German only.

August	1996	Intergenerational redistribution through the public sector – Methodology of generational accounting and its empirical application to Germany	Stephan Boll
August	1996	The impact of the exchange rate on Germany's balance of trade	Jörg Clostermann
October	1996	Alternative specifications of the German term structure and its information content regarding inflation	Sebastian T. Schich
November	1996	Enterprises' financing structure and their response to monetary policy stimuli An analysis based on the Deutsche Bundesbank's corporate balance sheet statistics	Elmar Stöss
January	1997	Reserve Requirements and Economic Stabilization	Ulrich Bindseil
June	1997	Direct investment and Germany as a business location	Thomas Jost
July	1997	Price Stability versus Low Inflation in Germany An Analysis of Costs and Benefits	Karl-Heinz Tödter Gerhard Ziebarth
October	1997	Estimating the German term structure	Sebastian T. Schich
October	1997	Inflation and Output in Germany: The Role of Inflation Expectations	Jürgen Reckwerth

February	1998	Problems of Inflation Measurement in Germany	Johannes Hoffmann
March	1998	Intertemporal Effects of Fiscal Policy in an RBC-Model	Günter Coenen
September	1998	Macroeconomic determinants of currency turbulences in emerging markets	Bernd Schnatz
January	1999	Die Geldmenge und ihre bilanziellen Gegenposten: Ein Vergleich zwischen wichtigen Ländern der Europäischen Währungsunion *	Dimut Lang
February	1999	Die Kapitalmarktzinsen in Deutschland und den USA: Wie eng ist der Zinsverbund? Eine Anwendung der multivariaten Kointegrationsanalyse *	Manfred Kremer
April	1999	Zur Diskussion über den Verbraucher- preisindex als Inflationsindikator – Beiträge zu einem Workshop in der Deutschen Bundesbank *	
July	1999	Monitoring Fiscal Adjustments in the European Union and EMU	Rolf Strauch
October	1999	Cyber money as a medium of exchange	Gabriele Kabelac

* Available in German only.

December	1999	Implicit Government Guarantees and Bank Herding Behavior	Rasmus Rüffer
December	1999	Implications of the new sesonal adjustment method Census X-12-ARIMA for current economic analysis in Germany	Robert Kirchner
February	2000	How Save Was the „Save Haven“? Financial Market Liquidity during the 1998 Turbulences	Christian Upper
May	2000	The determinants of the euro-dollar exchange rate – Synthetic fundamentals and a non-existing currency	Jörg Clostermann Bernd Schnatz