Analysing Divisia Aggregates for the Euro Area

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Abstract

In this paper, different Divisia monetary aggregates for the euro area are constructed over the period from 1980 to 2000. Theoretically, one main difference of these aggregates is their reaction to exchange-rate variations. Empirically, the aggregates are compared with respect to three issues. First, the demand for the Divisia aggregates is evaluated using the cointegrated VEC model and single-equation techniques, where stable demand functions are estimated. Second, the information content of these aggregates as regards future output is investigated. Evidence is presented that one of the Divisia aggregates has most information content from a forward-looking perspective. Third, using the P-star framework, the importance of money for future price movements is examined. Adapting an in-sample analysis, Divisia aggregates are important for HICP development and to some extent for GDP deflator movement. The out-of-sample forecasting exercise presents, on the one hand, evidence that simple-sum M3 includes more information for the HICP, whereas one of the Divisia aggregates helps to predict the future GDP deflator. On the other hand conspicuous control errors exist. In sum, the paper supports the view that money should have an important role in the conduct of monetary policy in the euro area.

Keywords: Divisia monetary aggregate; Money demand; Controllability; IS-curves; P-Star.

JEL Classification: E41, E52.

Zusammenfassung

In dieser Studie werden verschiedene Divisia-Aggregate für das Eurowährungsgebiet für den Zeitraum von 1980 bis 2000 berechnet. Ein wichtiger theoretischer Unterschied dieser Aggregate ist ihre Reaktion auf Wechselkursänderungen. Empirisch werden die Aggregate mit dem Summenaggregat M3 in Bezug auf drei Fragestellungen verglichen. Erstens werden Geldnachfragefunktionen von Divisia-Aggregaten mit Hilfe von kointegrierten VEC-Modellen und von Einzelgleichungsansätzen geschätzt. Es zeigt sich, dass stabile Gleichungen bestimmt werden. Zweitens wird der Informationsgehalt der Aggregate bezüglich der zukünftigen Outputentwicklung untersucht. Hierbei stellt sich heraus, dass eines der Divisia-Aggregate mehr Informationsgehalt als die anderen Aggregate besitzt. Drittens wird die Bedeutung des Geldes für die zukünftige Preisentwicklung analysiert. Bei der ex post-Analyse wird deutlich, dass Divisia-Aggregate die Entwicklung des Harmonisierten-Verbraucher-Preisindexes (HICP) und in geringerem Umfang die Entwicklung des BIP-Deflators beeinflussen. Die ex ante-Analyse verdeutlicht einerseits, dass das einfache Summenaggregate M3 mehr Informationsgehalt für die Entwicklung des HICP als die anderen Aggregate enthält, während eines der Divisia-Aggregate hilft die Vorausschätzung des zukünftigen DIP-Deflators zu verbessern. Andererseits gibt es erhebliche Kontrollfehler. Zusammenfassend unterstützt die Studie die Auffassung, dass die Geldmenge eine wichtige Rolle bei der Durchführung der Geldpolitik im Eurowährungsraum haben sollte.

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1 Introduction

The Eurosystem has the primary objective of maintaining price stability (see ECB, 2001). It organises its assessment of risks to price stability under two pillars. The first pillar gives money, especially M3, a prominent role, in line with the statement that inflation is a monetary phenomenon in the long-run, which is an essential principle of macroeconomic theory. The second pillar analyses a broad range of other economic and financial indicators relevant to future price development.

The monetary aggregate M3 is a simple-sum aggregate made up of different monetary components (see ECB, 1999). All the components included have the same weight and are considered to be perfect substitutes. The components that are excluded are assumed to have no substitutional relationship with money. Moreover, the theoretical foundation of this aggregation is weak. Therefore, Fase (2000), Spencer (1995) and Drake, Mullineux and Agung (1997), among others, have suggested constructing a Divisia monetary aggregate for the euro area. Divisia aggregates sum up the variable weighted growth rates of monetary components. This suggestion is adopted in the present study, where some difficulties have to be taken into account. The main problem is that of constructing the historical data. The euro area contains eleven (since January 2001 twelve) countries, which sample the national values of the different monetary components. Since January 1, 1999 exchange rates among the members of the Eurosystem have been irrevocably fixed. Before that date, exchange rates could change. The ECB (1999) suggests using the fixed exchange rates to combine national data for the euro

[1]Hans-Eggert Reimers, Hochschule Wismar, University of Technique, Business and Design, Postfach 12 10, D - 23952 Wismar, Germany, e-mail: h.reimers@wi.hs-wismar.de. Most of this research was conducted during my stay at the Deutsche Bundesbank as a visiting researcher (April-August 2001). The hospitality of the Bundesbank is greatly appreciated. I am grateful to seminar participants at the Deutsche Bundesbank for their useful comments. I wish to express my special thanks to Heinz Herrmann, Julian Reischle, Michael Scharnagl, Franz Seitz, and Karl-Heinz Tödter. The views expressed in this study are my own and not necessarily those of the Deutsche Bundesbank.

area data. The study sets alternative assumptions regarding the actions of economic agents to construct Divisia aggregates. Moreover, different exchange rate regimes are assumed, to calculate the aggregates. These settings result in one Divisia aggregate of national monetary components with fixed exchange rates, one Divisia aggregate of national monetary components with variable exchange rates and one aggregate of national Divisia aggregates, which are added up by accounting transaction cost weights (transaction cost weighted Divisia aggregate).

Despite the theoretical appeal of the Divisia aggregate, it is important to know its empirical properties. These properties are analysed with respect to three issues. First, the demand for the Divisia aggregates is evaluated using the cointegrated VEC model and single-equation techniques, where stable demand functions are estimated. Second, the information content of these aggregates as regards future output is investigated. For that purpose, IS-curves are estimated, which include, as additional regressors, money growth rates or money demand disequilibrium. In this study, evidence is presented that the transaction cost weighted Divisia aggregate has most information content from a forward-looking perspective. Third, using the P-star framework, the importance of money for future price movements is examined. Adapting an in-sample analysis, Divisia aggregates are important for the development of the harmonised index of consumer prices (HICP), and to some extent for GDP deflator movement. The out-of-sample forecasting exercise presents evidence that simple-sum M3 includes information for the future HICP, whereas one of the Divisia aggregates helps to estimate future GDP deflator. On the other hand, conspicuous control errors exist. In sum, no aggregate dominates the others regarding all analysed criteria.

The remainder is organised as follows. In the next section, the theoretical framework of multiplicative monetary aggregates is presented, with special emphasis on the effects of exchange rate variations. Section 3 contains the data and their descriptive analysis. Section 4 describes the money demand function investigation. Section 5 examines the importance of liquidity for the IS-relation and the link between prices and money. Section 6 analyses the information content of monetary aggregates for future price movements. Finally, Section 7 concludes.

2 Multiplicative monetary aggregates

2.1 General theory

Let us assume there is one economy. In this economy, there exists a representative agent. If his individual utility function is given as follows

$$u = u(c_1, c_2, l, m_1, m_2), \tag{1}$$

where c_1 and c_2 are consumer goods, l is leisure, and m_1 and m_2 are financial assets with a potential for moneyness, then weak separability implies that some arguments of the utility function can be put together. This is possible if the marginal rate of substitution between any two goods of the same group is independent of the quantity of goods in another group. On the assumption of weak separability for the two financial assets, the utility function may be written as

$$u = u(c_1, c_2, l, M(m_1, m_2)) \tag{2}$$

with

$$\partial \frac{(\partial m_1/\partial m_2)}{\partial c_i} = 0 \quad \text{for} \quad i = 1, 2. \tag{3}$$

The marginal rate of substitution between the financial assets m_1 and m_2 is not influenced by changing quantities of c_1. Weak separability is the necessary condition for generating the structure of a utility-tree (see Reischle, 2000, pp. 184-217). The total utility function is a function of sub-utility functions

$$u = f(u_c(C), u_l(l), u_m(M)). \tag{4}$$

With utility levels u_c and u_l given, utility maximisation will be reduced to the maximisation of u_m under the constraint

$$\sum_{i=1}^{2} p_i m_i = y_m, \tag{5}$$

where p_i is the price and m_i the quantity of the financial asset i, y_m is the expenditure on M. The demand for the particular components of M depends only on the relative prices (p_m) of the particular financial assets and on the amount of expenditure spent on financial assets

$$m_i = \theta_i(p_m, y_m) \quad \text{for} \quad i = 1, 2. \tag{6}$$

The total income $y = y_c + y_l + y_m$ and the prices p_l and p_c affect the demand for group m assets only via y_m (general substitution effect). When y_m is given, p_c and p_l can be disregarded. All prices p_c exert a proportionate influence on m_i.

An alternative way of dealing with these problems is to construct Divisia aggregates, as proposed by Barnett (1978, 1980). Let us assume that there is a benchmark asset with yield R_t, which provides no monetary services and is held solely to transfer wealth intertemporally. Holding the liquid asset i with yield r_{it} costs $R_t - r_{it}$ per unit of currency in period t. Total transaction costs in period t can be expressed as

$$K_t = \sum_{i=1}^{L} (R_t - r_{it}) m_{it}, \tag{7}$$

where m_{it} is the value of monetary component i and L is the number of considered components. The expenditure share of the i^{th} asset is

$$s_{it} = \frac{(R_t - r_{it}) m_{it}}{K_t} = \frac{(R_t - r_{it}) m_{it}}{\sum_{i=1}^{L} (R_t - r_{it}) m_{it}} \,. \tag{8}$$

Real user costs are

$$\pi_{it} = \frac{R_t - r_{it}}{1 + R_t}. \tag{9}$$

Furthermore, let us assume that the transaction technology can be described by the general, twice differential, homogeneous function

$$m_t = M(m_{it}, \cdots, m_{Lt}) \,. \tag{10}$$

Minimizing of transaction costs (7), subject to (10), results in a Divisia monetary index

$$d \ln DM_t = \sum_{i=1}^{L} s_{it} d \ln m_{it} \,, \tag{11}$$

where $d\ln$ denotes the ln-differential of a variable. In discrete time, usually the Törnquist-Theil approximation of the Divisia index is used

$$\Delta \ln DM_t = \sum_{i=1}^{L} \tilde{s}_{it} \Delta \ln m_{it} \tag{12}$$

with weights $\tilde{s}_{it} = \frac{1}{2}(s_{it} + s_{i,t-1})$ (see Barnett, Offenbacher and Spindt, 1984, p. 1052). The price dual of the Divisia quantity index is given by

$$\Delta \ln Pd_t = \sum_{i=1}^{L} \tilde{s}_{it} \Delta \ln(R_t - r_{it}) \,. \tag{13}$$

Equivalently, it is calculated by

$$Pd_t = \frac{\sum_{i=1}^{L}(R_t - r_{it})m_{it}/(1+R_t)}{DM_t}$$

since $Pd_t \cdot DM_t = K_t$.

It is worth noting that the Divisia index refers to the growth of monetary services provided by the monetary components (see Gaab and Mullineux, 1996). The levels of monetary services have to be recovered following normalisation. The user cost s_i can be regarded as the cost of purchasing an additional unit of monetary service of the i-th monetary component. A disadvantage of the Divisia aggregate is that it measures money on the base of the changes in the logarithm of its components. It can not handle the introduction of new assets. Because the logarithm of zero is minus infinity, the formula for the Divisia aggregate implies that the growth rate of the Divisia index equals infinity when a new asset is introduced. Thus, in a period when a new monetary asset is introduced, one has to set the growth rate of the new asset to zero. Gaab and Mullineux (1996) mention further problems posed by calculating Divisia indexes.

So far, the user costs have been determined without risk or on the assumption of the risk neutrality of the consumer. The inclusion of uncertainty changes the utility function. Barnett, Hinich and Yue (2000) assume that the utility function is

$$u = E_t \sum_{t=0}^{\infty} \beta^t u(c_1, c_2, l, M(m_1, m_2)),$$

where E_t is the expectation operator and β the discount factor. On the assumption of risk neutrality, the user costs change to

$$\pi_{it}^{e} = \frac{E_t(R_t - r_{it})}{E_t(1+R_t)}$$

The expected interest rates are included in this definition. On the assumption of risk aversion, Barnett, Liu and Jensen (2000) show that the user costs include a risk premium ϕ_{it}

$$\pi_{it}^{g} = \pi_{it}^{e} + \phi_{it}.$$

Barnett, Liu and Jensen (2000) show that this relationship is approximated to by

$$\pi_{it}^{g} = \frac{E_t(R_t) - E_t(r_{it} - HCov(r_{it}, \Delta c))}{1 + E_t(R_t)},$$

where H is the Arrow-Pratt measure of relative risk aversion and Cov is the covariance between yield r_{it} and the growth rate of consumption Δc. H is defined as $H = -c\frac{u''}{u'}$ where u' (u'') is the first (second) derivative of the utility function.

2.2 Euro area aggregates

There exist a few approaches to determining a Divisia aggregate for the euro area. They differ regarding the assumptions about the representative agent.

Assumption of one representative agent: At first, it is assumed that there is one representative agent for the whole euro area. This agent has one benchmark interest rate, which is the highest rate among all relevant national interest rates. Following the aggregation proposal of the ECB (1999), fixed exchange rates are used to construct the euro area historical data. In this sense, (11) is applied to all relevant components of the individual euro area countries' monetary aggregate:

$$\Delta \ln DM_t^1 = \sum_{i=1}^{L} \sum_{j=1}^{J} \tilde{s}_{ijt} \Delta \ln m_{ijt} \bar{e}_j, \tag{14}$$

where $\tilde{s}_{ij}$ ($m_{ij}\bar{e}_j$) is the i-th expenditure share (component) of the j-th euro area member and J the number of euro area members. It is worth noting that the irrevocably fixed conversion rates of 31 December 1998 ($\bar{e}_j$) are applied to construct the expenditure shares and monetary components.

Because not all countries deliver historical data for the components, Stracca (2001a) suggests using the M3 components of the euro area and aggregate interest rate series to construct an aggregate:

$$\Delta \ln DM_t^2 = \sum_{i=1}^{L} \tilde{s}_{it}^{euro} \Delta \ln m_{it}^{euro} \tag{15}$$

where $m_{it}^{euro} = \sum_{j=1}^{J} m_{ijt}\bar{e}_j$, applying fixed exchange rates. The aggregate interest rate ($\bar{r}_{it}$) is determined by GDP weights $\bar{r}_{it} = \sum_{j=1}^{J} w_j^{GDP} r_{ijt}$. It is worth noting that DM^1 equals DM^2 if $r_{i1t} = r_{i2t} = \cdots = r_{iJt}$ for $i = 1, \cdots, L$ and R_t in (14) is identical to R_t in (15).

However, both approaches have in common the fixed exchange rate assumption. This assumption is at odds with historical experience. Therefore, Wesche (1997) assumes one representative agent who accounts for variations in exchange rates. Constructing a European monetary aggregate, it is assumed that consumers hold a diversified portfolio of European currencies with different degrees of liquidity (see Wesche, 1997). The stock of monetary assets is redefined to account for currencies of different denominations. This means that the representative consumer is assumed to hold monetary assets, denominated in different European currencies $m_{ijt}e_{jt}$, where m_{ijt} is the i-th monetary asset denominated in the j-th country's currency and e_{jt} is the j-th country's exchange rate, relative to a weighted currency basket like the Ecu.

In addition, the own rate of return r_{it} of a component monetary asset has to take into account the expected depreciation or appreciation of the respective currency relative to the Ecu. The user cost for the European Divisia index thus becomes

$$\pi_{ijt}^{e} = \frac{E_t(R_t - (r_{ijt} + \psi_{jt}))}{E_t(1 + R_t)} \tag{16}$$

with

$$E_t \psi_{jt} = \frac{e_{jt+1}^{e} - e_{jt}}{e_j}$$

being the expected depreciation of the jth country's currency and

$$E_t R_t = \max(E_t(R_{jt} + \psi_{jt})) \quad \text{for } j = 1, \cdots, J \tag{17}$$

being the European benchmark yield, which is the highest yield on a portfolio of European bonds, corrected for the expected depreciation of the exchange rate. The Divisia aggregate becomes

$$\Delta \ln DM_t^3 = \sum_{i=1}^{L} \sum_{j=1}^{J} \bar{\bar{s}}_{ijt} \Delta \ln m_{ijt} e_{jt}, \tag{18}$$

where $\bar{\bar{s}}_{ijt}$ involves π_{ijt}^{e}. Without variations in the exchange rates, DM^3 equals DM^1. Equation (16) may be further simplified if the uncovered interest rate parity holds. In this case, the different national interest rates (foreign interest rates) of one component, except for one country, are substituted by the interest rate of one country (home country). It should be stressed that a common characteristic of the three proposals is that they do not account for differences in national behaviour and national financial systems.

Assumption of representative national agents: The alternative is that there are country-specific agents who determine a national monetary aggregate and that, in the second step, these national series are aggregated. For example, the Divisia aggregates DM_{jt} denominated in national currencies, are summed up

$$DM_t^4 = \sum_{j=1}^{J} DM_{jt} e_{jt}, \tag{19}$$

where the national Divisia indices are normalised in such a way that they are equal to the corresponding simple-sum aggregate. In this sense, DM_t^4 is comparable to a simple-sum aggregate for the euro area. Bayoumi and Kenen (1993) criticise an additive aggregation of the levels, since it neglects the fact that differences in behaviour may cause members of the euro area to use money at different intensities. They propose summing up the weighted growth rates of national monetary aggregates.

$$\Delta \ln DM_t^5 = \sum_{j=1}^{J} w_j^{GDP} \Delta \ln DM_{jt}. \tag{20}$$

The weights w_j are determined by constant GDP shares. Beyer, Doornik and Hendry (2001) mention that (20) is distorted if the GDP shares and money shares differ considerably. Therefore Beyer et al. (2001) suggest using variable money shares

$$\Delta \ln m_t = \sum_{j=1}^{J} w_{jt}^{Mon} \Delta \ln m_{jt}, \tag{21}$$

where w_{jt}^{Mon} is recursively determined. If this procedure is adopted for a Divisia aggregate, it is

$$\Delta \ln DM_t^6 = \sum_{j=1}^{J} w_{jt}^{Mon} \Delta \ln DM_{jt}, \tag{22}$$

where $w_{jt}^{Mon} = \frac{DM_{jt}e_{jt}}{\sum_{j=1}^{J} DM_{jt}e_{jt}}$ gives the weights.

Since the weights of the Divisia aggregate result from minimising transaction costs for a given transaction technology, it seems sensible to construct weights depending on expenditure shares, as proposed by Reimers and Tödter (1994). The euro area transaction costs are

$$K_t^{euro} = \sum_{j=1}^{J} K_{jt} e_{jt}.$$

The national expenditure shares are given by

$$w_{jt}^K = \frac{K_{jt} e_{jt}}{K_t^{euro}}.$$

Hence the euro area Divisia aggregate is

$$\Delta \ln DM_t^7 = \sum_{j=1}^{J} w_{jt-1}^K \ln \Delta DM_{jt}. \tag{23}$$

This aggregate accounts for differences in national financial systems. If the national benchmarks converge to one value and the national interest rates of the components converge to specific values, it is identical to an aggregate where the components are summed up and afterwards a Divisia aggregate is calculated.

Analysis of the exchange rate effect: One main difference between these seven aggregates is the assumptions regarding the exchange rates. Therefore the effects of variations in the exchange rate differ. Following Beyer et al. (2001), it is easy to show that the exchange rate effects are small by comparing DM^7 with DM^4. For the additive aggregate, it is

$$\frac{\partial DM^4}{\partial e_{js}} = DM_{js}. \tag{24}$$

For the growth-rate aggregation of DM^3, it is

$$DM_t^3 = \exp\left(\sum_{v=1}^{t}\sum_{i=1}^{L}\sum_{j=1}^{J} \bar{\tilde{s}}_{ijv} \Delta \ln m_{ijv} e_{iv-1}\right).$$

The partial derivative with respect to a change in e_{jt} at $t = s + 1$ is:

$$\begin{aligned} \frac{\partial DM_t^3}{\partial e_{js}}_{|t=s+1} &= \frac{\partial}{\partial e_{js}} \left(\sum_{v=1}^{t}\sum_{i=1}^{L}\sum_{j=1}^{J} \bar{\tilde{s}}_{ijv} \Delta \ln m_{ijv} e_{iv-1}\right)_{|t=s+1} DM_t^3 \\ &= \frac{\partial}{\partial e_{js}} \left(\sum_{i=1}^{L}\sum_{j=1}^{J} \bar{\tilde{s}}_{ijs} \Delta \ln m_{ijs+1} e_{js}\right) DM_t^3 \\ &= \sum_{i=1}^{L}\sum_{j=1}^{J} \left(\frac{\partial \bar{\tilde{s}}_{ijs}}{\partial e_{js}} \Delta \ln m_{ijs+1} e_{js} + \bar{\tilde{s}}_{ijs} \Delta \ln m_{ijs+1}\right) DM_t^3, \end{aligned} \tag{25}$$

where

$$\begin{aligned} \frac{\partial \bar{\tilde{s}}_{ijs}}{\partial e_{js}} &= \frac{\partial}{\partial e_{js}} \frac{1}{2} \left(\frac{(R_s - r_{ijs} - \psi_{js}) m_{ijs} e_{js}}{\sum_{i=1}^{L} (R_s - r_{ijs} - \psi_{js}) m_{ijs} e_{js}} \right. \\ &\quad \left. + \frac{(R_{s-1} - r_{ijs-1} - \psi_{js-1}) m_{ijs-1} e_{js-1}}{\sum_{i=1}^{L} (R_{s-1} - r_{ijs-1} - \psi_{js-1}) m_{ijs-1} e_{js-1}} \right) \\ &= \frac{\partial}{\partial e_{js}} \frac{1}{2} \left(\frac{(R_s - r_{ijs} - \psi_{js}) m_{ijs} e_{js}}{\sum_{i=1}^{L} (R_s - r_{ijs} - \psi_{js}) m_{ijs} e_{js}} \right) = 0, \end{aligned}$$

if R_s is an interest rate of country j or if $\frac{\partial \psi_{js}}{\partial e_{js}} = 0$, than the expected exchange rate is independent of variations in the actual exchange rate. These conditions imply that the expenditure shares do not vary with the exchange rates. This simplifies equation (25). Hence,

$$\frac{\partial DM_t^3}{\partial e_{js}}_{|t=s+1} = \sum_{i=1}^{L} \left(\bar{\tilde{s}}_{ijs} \Delta \ln m_{ijs+1}\right) DM_{s+1}^3 = \Delta \ln DM_{js} \cdot DM_{s+1}^3. \tag{26}$$

For the growth-rate aggregation of DM^7 it is

$$DM_t^7 = \exp\left(\sum_{v=1}^{t}\sum_{j=1}^{J}\frac{K_{jv-1}e_{jv-1}}{\sum_{i=1}^{J}K_{iv-1}e_{iv-1}}\Delta\ln DM_{jv}\right)$$

The partial derivative with respect to a change in e_{jt} at $t = s+1$ is

$$\begin{aligned}
\frac{\partial DM_t^7}{\partial e_{js}}_{|t=s+1} &= \frac{\partial}{\partial e_{js}}\left(\sum_{v=1}^{t}\sum_{j=1}^{J}\frac{K_{jv-1}e_{jv-1}}{\sum_{i=1}^{J}K_{iv-1}e_{iv-1}}\Delta\ln DM_{jv}\right)_{|t=s+1} DM_t^7 \\
&= \frac{\partial}{\partial e_{js}}\left(\sum_{j=1}^{J}\frac{K_{js}e_{js}}{\sum_{i=1}^{J}K_{is}e_{is}}\Delta\ln DM_{js+1}\right)DM_t^7 \\
&= \left(\Delta\ln DM_{js+1}^7 - \frac{\sum_{j=1}^{J}K_{js}e_{js}\Delta\ln DM_{js+1}^7}{\sum_{i=1}^{J}K_{is}e_{is}}\right)\frac{K_{js}}{\sum_{i=1}^{J}K_{is}e_{is}}DM_t^7 \\
&\simeq (\Delta\ln DM_{js+1}^7 - \Delta\overline{\ln DM^7}_{s+1})DM_t^4, \qquad (27)
\end{aligned}$$

where

$$\Delta\overline{\ln DM^7}_{s+1} = \frac{\sum_{j=1}^{J}K_{js}e_{js}\Delta\ln DM_{js+1}^7}{\sum_{i=1}^{J}K_{is}e_{is}}$$

is the average growth rate and

$$\frac{K_{js}}{\sum_{i=1}^{J}K_{is}e_{is}}DM_t^7 \simeq DM_t^4.$$

The exchange rate effect in (27) is of a smaller order than in (24) unless $(\Delta\ln DM_{js+1}^7 - \Delta\overline{\ln DM^7}_{s+1}) \simeq 1$. It is worth noting that DM^1 and DM^2 do not react to exchange rate variations. As far as the actual exchange rate e_{it} diverges from the fixed exchange rate $\bar{e}_i$, the DM^1 and DM^2 are biased compared with aggregates that are constructed using variable exchange rates.

3 Data

In this study, data from 1980 through 2000 are used. As a measure of M3, quarterly averages of the month-end stocks of M3 are used (Source: ECB, in billions of euro, using the definition of April 2000). The main components of M3 are currency in circulation, overnight deposits, deposits with an agreed maturity of up to two years, deposits redeemable at notice of up to three months, repurchase agreements, debt securities issued with a maturity of up to two years and money market fund shares/units and money market paper (see Table 1). The Bundesbank has monthly data on seven categories for five countries (Germany, France, Spain, Portugal and Finland) and for the whole euro area. Overnight deposits are constructed using M1H from the Bundesbank converted into euro via the irrevocable fixed conversion rates of 31 December 1998. The attempt to do the same for time and saving deposits, using M3H, was not successful. Therefore a block is constructed, representing the stocks of Austria, Italy, Belgium, Netherlands, Luxembourg and Ireland.

A key item of information necessary to derive Divisia monetary aggregates is the own rates of return on the monetary components. For this purpose, it is necessary to estimate series of rates of return over the sample period 1980Q1-2000Q4. The construction is split into two parts. From 1980 till 1997 country-specific data are collected. Since 1998 euro area data have been used. They are published by the ECB in its Monthly Bulletin (Table 2.6: Money market interest rates; Table 2.9: Retail bank interest rates, deposit interest rates).

Data collection before 1998 is more complicated. The ECB publishes the retail interest rates of the member countries. Following Dedola, Gaiotti and Silipo (2001), in some cases the information is completed by data from national sources. They are taken from the database of the BIS or IMF. The central bank interest rates, money market rates and some public bond yields are from International Financial Statistics (IFS). Non-available data points are replaced by linear approximations of the neighbouring data points. To determine the corresponding interest rates of the block components, the country weights of the monetary component are calculated for the period 1998 to 2000. These weights are used to generate the composite interest rates of the block components. M3 country weights are used to determine the euro area central bank interest rate (i_cen), public bond yields (R^{bo}) and the money market rate (R^{mo}). Quarterly data are calculated as the average of three monthly observations. Moreover, the own interest rate of M3 (R^{M3}) is taken from Calza, Gerdesmeier and Levy (2001).

Nominal and real GDP from 1991Q1 is calculated on the basis of the ESA95 System of Accounts (Deutsche Bundesbank). Using the data of Stracca (2001b), the series are supplemented by linking their growth rates backwards until 1980Q1. The price index is the implicit GDP deflator. Alternatively, the HICP is used. It is available from 1991Q1 onwards. Collecting the CPI data of the euro area countries and determining

Table 1: Monetary components of M3 and corresponding interest rates of the euro area

Monetary component	Own rate of return
Currency in circulation (BG)	Zero
Overnight deposits (SE)	Interest rate of overnight deposits
Deposits with an agreed maturity of up to two years (time deposits, TE)	Time deposit rate up to 1 year
Deposits redeemable at notice up to three months (savings deposits, SP)	Savings deposit rate up to 3 months
Repurchase agreements (RE)	3-month money market rate
Money market fund shares/units and money market paper (MM)	3-month money market rate
Debt securities issued with a maturity of up to two years (BS)	12-month money market rate

variable GDP weights allows us to construct the series backwards until 1980Q1.

Weak separability requires that the empirical data can be described by a "well-behaved" utility function, i.e. individuals reveal no preferences inconsistent with the generalised axiom of revealed preference (GARP). An aggregate satisfies GARP if

$$p^j m^j \geq p^j m^i \tag{28}$$

applies, and at the same time

$$p^i m^i > p^i m^j \tag{29}$$

fails to apply (see Varian, 1982 and 1983). m^i and m^j are vectors of financial assets, p^i and p^j are corresponding prices. If i and j are interpreted as time indices, then p^i, p^j, m^i and m^j can be interpreted as combinations of prices and quantities in two different periods. If condition (28) holds, m^j is chosen although combination m^i would be cheaper. In this case, individuals reveal a preference for m^j. In contrast, if condition (29) holds, individuals prefer m^i to m^j. If both conditions are valid, there is a contradiction that cannot be represented by a well-behaved utility function. Hence the pairwise comparisons allow us to test the necessary condition for the weak separability of a utility function (for an extended discussion of GARP tests see Reischle, 2000, pp. 274-309).

Table 1 shows the quantities m^i. Real user costs are interpreted as prices p^i. Following Barnett (1978) the real user costs $\pi_{i,t}$ of financial asset i are defined in equation

(9). The data are available for 1997M12 - 2000M12, 37 monthly observations.

When constructing a Divisia index, one has to select a benchmark asset. As mentioned above, it should be the rate of return on a capital certain financial asset providing no monetary services. However, "pure" examples of such benchmark assets are hardly available in practice. The long-term government bond yield with a maturity of 10 years for the euro area is therefore used as a convenient proxy.

Table 2: Results of tests of GARP

Elements of the utility function	Elements of the sub-utility function	Nominal values	Real values
BG, SE, TE, SP, GR, BS		0	0
BG, SE, TE, SP, GR, BS	BG, SE, SP, TE, BS	0	0
BG, SE, TE, SP, GR, BS	BG, SE, SP, TE, GR	1	1
BG, SE, TE, SP, GR, BS	BG, SE, SP, TE (= M2)	1	1
BG, SE, TE, SP, GR, BS	BG, SE, SP	1	1
BG, SE, TE, SP, GR, BS	BG, SE, TE	1	1
BG, SE, TE, SP, GR, BS	BG, SE (= M1)	0	0

To determine the real values, the HICP is used. Abbreviations are defined in Table 1. GR = MM + RE.

Concerning this sample, no violations of GARP can be observed for the whole aggregate (see Table 2). This result is in line with Scharnagl (1996) for Germany. When particular components within this group are summed up, weak separability can be shown for M1 and M2+BS. For M2, this property must be rejected, as in this case GARP does not hold. There is no difference between the nominal and real values.

To reduce the complexity of the study, only the aggregates DM^1, DM^3 and DM^7 are analysed. DM^2 is investigated by Stracca (2001a). The calculation of DM^3 needs values for the expected exchange rate. Different proposals determining the series exist. On the assumption of perfect foresight, the expected change equals the current change. One disadvantage of this procedure is that the resulting series are very volatile. Furthermore, the depreciation or appreciation rates obtained do not isolate possible risk premia in a currency. Assuming that the purchase power parity (PPP) holds, PPP exchange rates may be an alternative to calculating expected exchange rates. Since different suggestions of PPP exchange rates exist, a statistical method is used to deter-

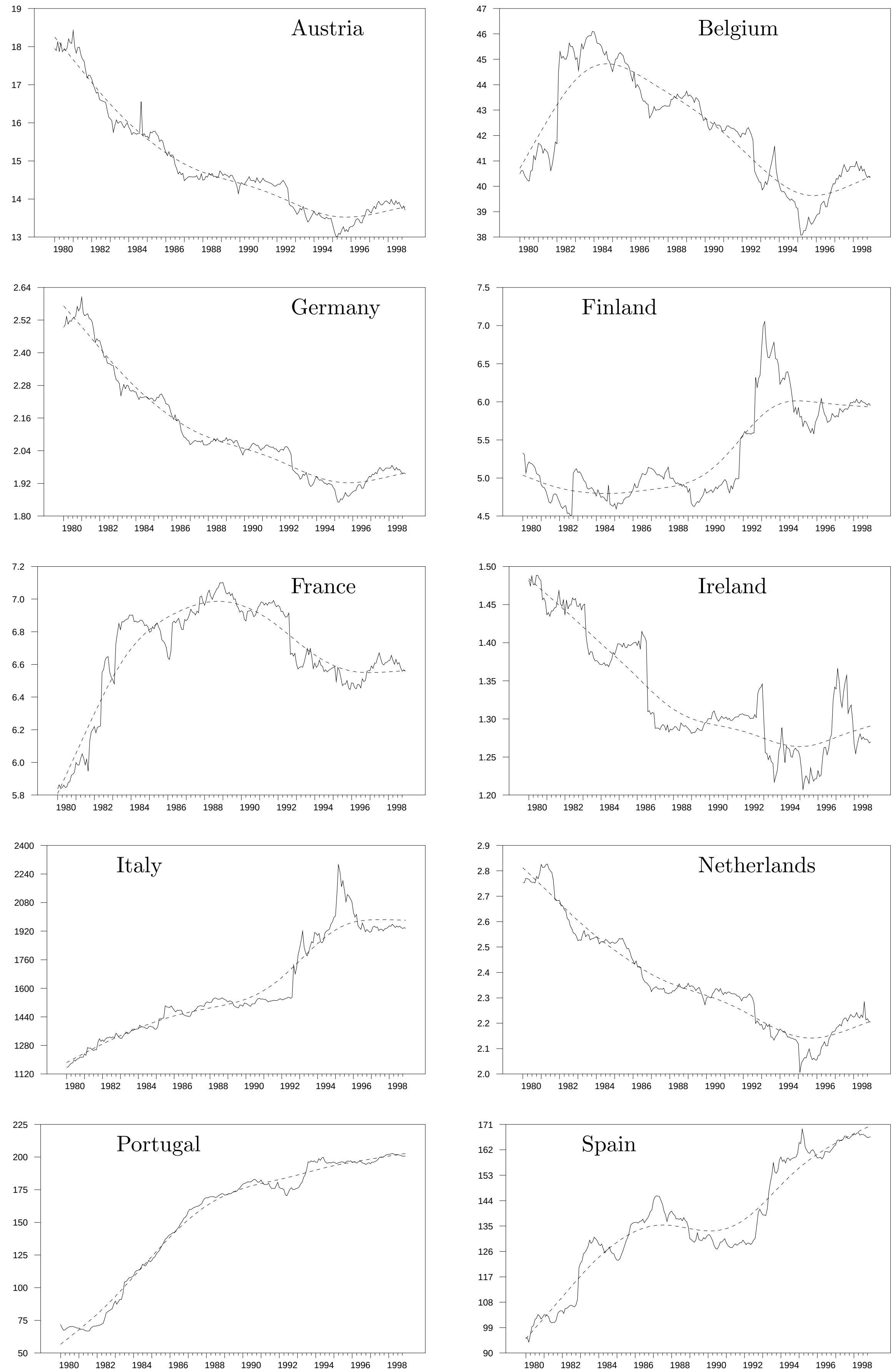

Figure 1: Exchange rates against Ecu (solid lines) and their trends (dashed lines) determined by Hodrick-Prescott filter. 1980M1 - 1998M12.

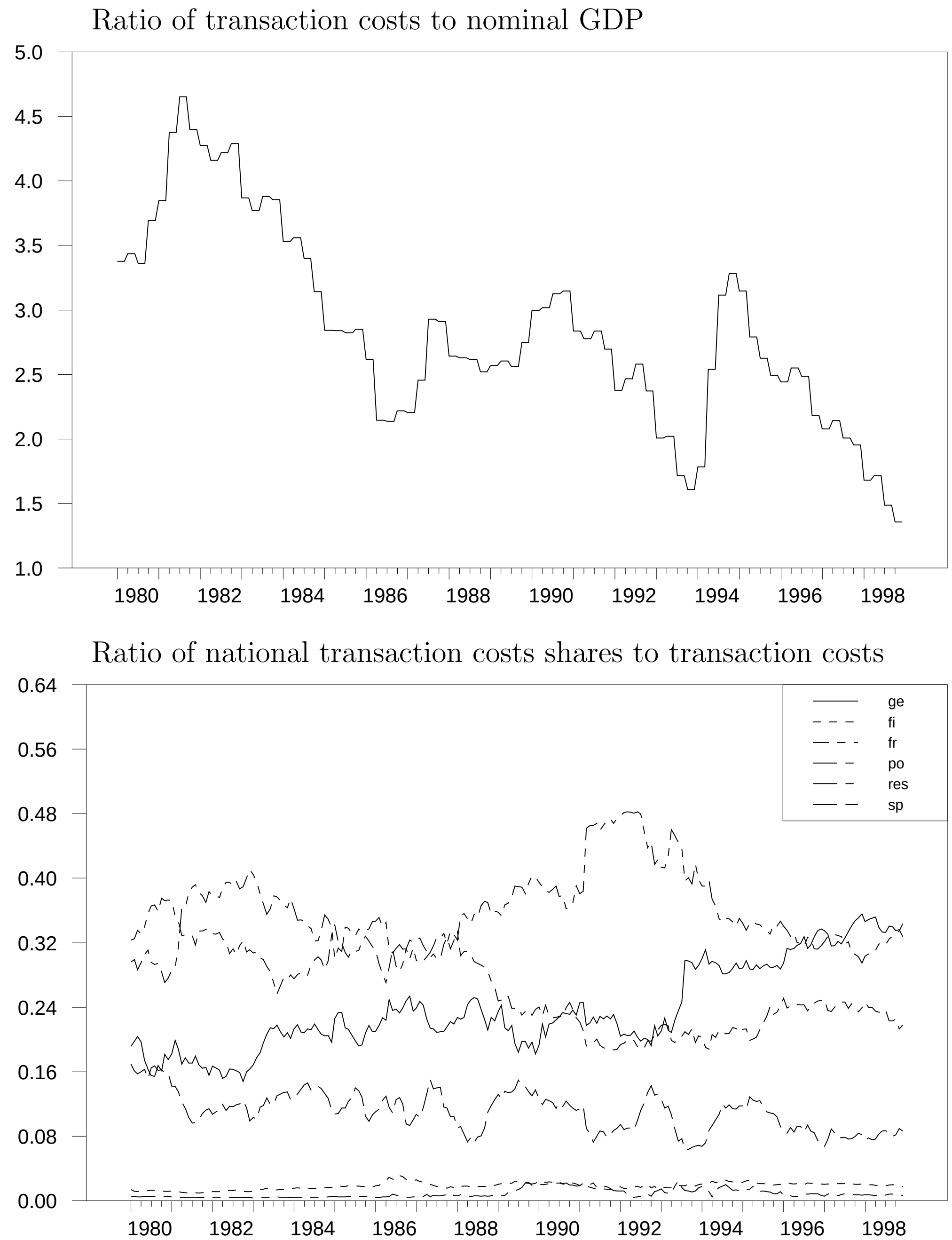

Figure 2: Ratio of transaction costs to nominal GDP in per cent (upper panel) and ratio of national transaction costs shares to whole transaction costs (lower panel), where the following abbreviations are used: ge (Germany), fi (Finland), fr (France), po (Portugal), sp (Spain), res (rest of the euro area), 1980-1998.

mine the expected exchange rates. In this study, the Hodrick-Prescott filter is used (see Appendix). National exchange rates against the euro and the filtered series are given in Figure 1. The implied depreciation or appreciation rates are variable but smoothed. The rates are lower at the end of the nineties than at the beginning of the eighties.

DM[7] gives hints of euro area transaction costs, which are determined for the national aggregates and summed up using the current ecu exchange rates. Its quarterly values are set in relation to nominal GDP. Figure 2 gives the ratio of the transaction costs to nominal GDP in percentage terms. In general, the transaction costs decline. One reason is that the benchmark interest rate declines. Another is that the own interest rates of monetary components increase. In some cases they move in the direction of benchmark interest rate. In Figure 2, the interest rate cycles are clearly apparent.

Figure 2 also presents national transaction costs relative to euro area transaction costs. The share of Germany increases owing to the larger share of currency in circulation, whereas the shares of Spain and France decline. This may result from the decrease in the benchmark in those countries. Some statistics of the transaction cost shares are exhibited in Table 3. It is worth noting that the transaction cost shares diverge remarkable from the M3 shares, and from the GDP weights. Furthermore, it is noteworthy that the transaction cost shares are not stationary. An augmented Dickey-Fuller-test as well as a Phillips-Perron-test indicate that the null hypothesis of one unit root in the series is not rejected.

Table 3: Descriptive statistics for national transaction cost shares

Country	Germany	Finland	Portugal	Spain	France	others
Mean of shares	0.237	0.018	0.009	0.111	0.264	0.360
Value 1997M12	0.356	0.021	0.007	0.082	0.239	0.295
Mean of M3 shares	0.284	0.016	0.013	0.087	0.237	0.364
Value 1997M12	0.281	0.016	0.022	0.112	0.204	0.364
Mean of GDP weights	0.310	0.018	0.016	0.096	0.244	0.316
Value 1997Q4	0.329	0.018	0.017	0.103	0.229	0.304

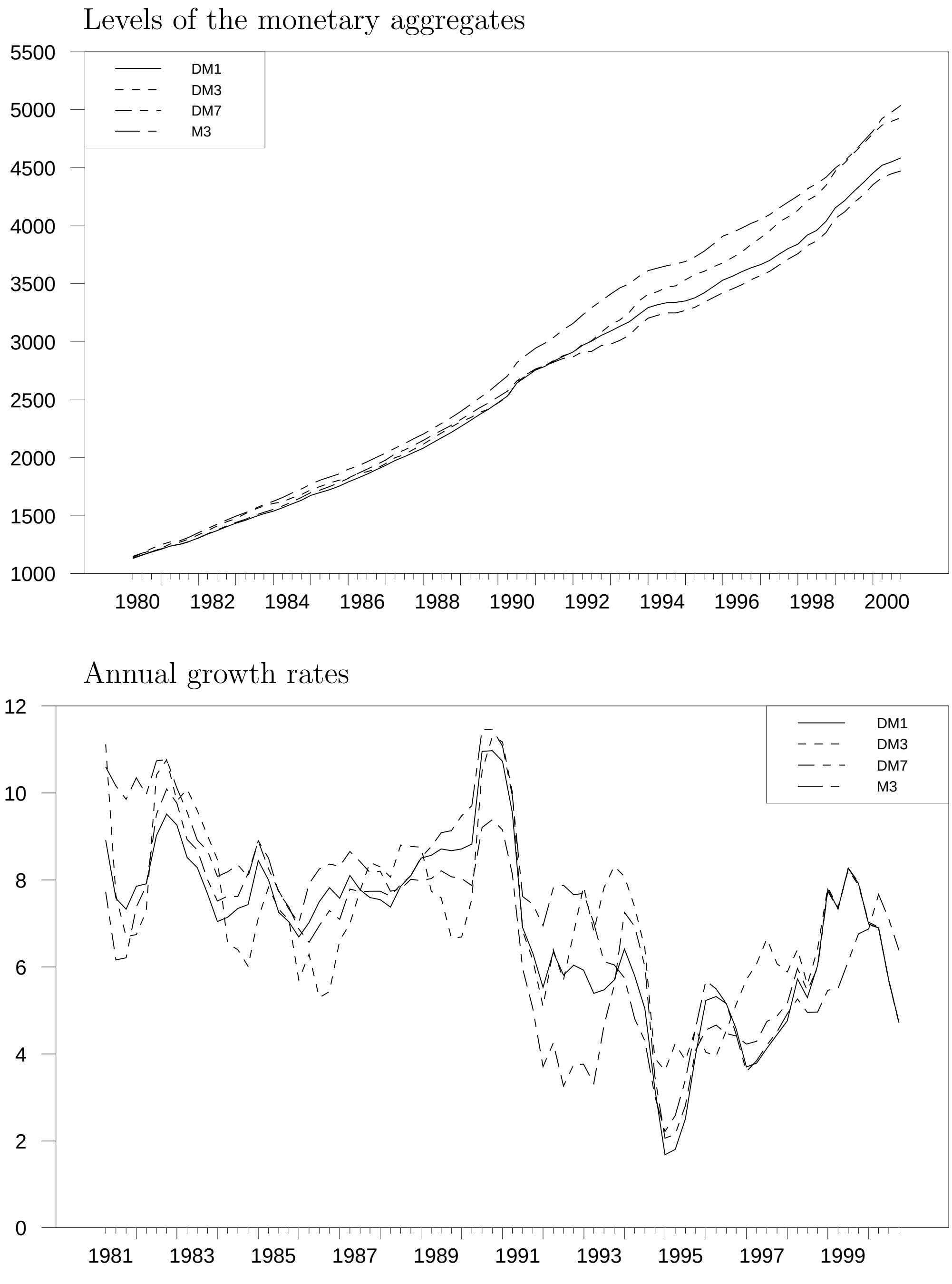

Figure 3: Levels of the different monetary aggregates in euro billions, 1980-2000 (upper panel); Annual growth rates of the different monetary aggregates in per cent, 1981-2000 (lower panel).

The multiplicative aggregates are individually constructed for the sample period 1980 to 1997. For the period 1998 to 2000, the existence of fixed exchange rates is assumed and a Divisia aggregate is calculated for the whole euro area. These values are used to complete the individually constructed series. The development of the monetary aggregates is given in Figure 3. They are seasonally adjusted using X12-ARIMA routine of EVIEWS4.0 (multiplicative). They are normalised in such a way that their values are identical in the second month of 1980. It is apparent that the level values of the multiplicative aggregates are smaller at the end of the sample period than official M3. All aggregates reflect German unification in the middle of 1990. Looking at the annual growth rates, the differences in the series are more pronounced (see Figure 3, lower panel). The descriptive test statistics are given in Table 4. The average annual growth rate of M3 and its volatility are higher than the growth rates of the other aggregates and their volatility. The correlation is strong among $\Delta_4 \ln M3$ and $\Delta_4 \ln DM^1$ as well as $\Delta_4 \ln DM^1$ and $\Delta_4 \ln DM^7$. These results indicate that the aggregates may cover the same long-run movement, however, may exhibit small but important differences in the short-term development.

Table 4: Descriptive statistics of annual growth rates of M3 and Divisia M3 (DM^1, DM^3 and DM^7)

Statistic	Δ_4 ln M3	$\Delta_4 \ln DM^1$	$\Delta_4 \ln DM^3$	$\Delta_4 \ln DM^7$
Mean	0.073	0.068	0.071	0.067
Maximum	0.115	0.110	0.114	0.101
Minimum	0.022	0.017	0.036	0.021
Std. Dev.	0.022	0.019	0.018	0.020
J.B.	1.534	2.471	1.555	5.836
	(.464)	(.291)	(.460)	(.054)
Correlation with Δ_4 ln M3		0.898	0.731	.718
Correlation with Δ_4 ln DM^1			0.800	0.900
Correlation with Δ_4 ln DM^3				0.728

J.B.: Jarque-Bera-test of normality, its p-value in parentheses. The information period is 1981Q2 - 2000Q4. Variables are seasonally adjusted.

4 Money demand systems and controllability

According to the theory presented demand for the Divisia aggregates should depend positively on total expenditure and negatively on Divisia price duals (pd). Total expenditure is approximated by euro area GDP (y). The long-run demand for log real Divisia (dm) is specified as follows

$$dm_t = \beta_0 + \beta_1 y_t + \beta_2 pd_t + e_t,$$

where e_t is a stationary process. This equation is more restricted than the specification by Stracca (2001a). His equation includes a squared term of pd_t. On the assumption that pd_t is an I(1)-process, then pd_t^2 is not an I(1)-process. This would enormously complicate the analysis.

Divisia price duals (see equation 13) are assumed to represent the opportunity cost of money holding. It depends on own interest rates and the benchmark interest rate. To test the controllability of money demand by central bank interest rates, i_cen are additionally included. For example, Johansen and Juselius (2001) mention the importance of controllability for monetary policy. Referring to central banks their main instruments are central bank interest rates. On the assumption that the central bank conducts monetary policy with a money growth target, a convincing policy presupposes that the target is controllable by the central bank. Johansen and Juselius (2001) account for the nonstationarity and cointegrating properties of the considered variables and define controllability by a condition on the elements of a long-run impact matrix Θ, which is determined by the orthogonal complements of the cointegrating matrix C and the loading matrix B (see equation 30). Therefore, a stationary variable which is a linear combination of $C'x_t$ cannot be controlled by this rule. In the simple case of one target and one instrument, Johansen and Juselius (2001) show that the long-run impact of a shock (an intervention) to the instrument variable is bound to affect the target variable. Controllability is inconsistent with long-run neutrality of target to instrument. To answer this controllability question, the systems analysed include a real Divisia aggregate, real GDP, price dual and central bank interest rates.

The starting point of the empirical analysis is a vector autoregressive (VAR) model of the lag order p

$$x_t = \nu + A_1 x_{t-1} + \cdots + A_p x_{t-p} + \epsilon_t$$

where ϵ_t is the white noise process and x_t a K-dimensional nonstationary process. Assuming that the integrating order of the variables is at most one and that the variables are cointegrated, the VAR-model may be reparametrised as a vector error correction model.

$$\Delta x_t = \nu + \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{p-1} \Delta x_{t-p+1} + \Pi x_{t-p} + \epsilon_t$$

If Π has a cointegrating rank of r it may be rewritten as $\Pi = BC$ where B (C') are $K \times r$-matrices of rank r. The Johansen-procedure allows us to test the cointegrating space and gives maximum likelihood estimates of the unknown coefficient matrices (see Johansen, 1988, 1991). The impact matrix is defined as:

$$\Theta = C'_\perp (B'_\perp (I - \sum_{i=1}^{p-1} \Gamma_i) C'_\perp)^{-1} B'_\perp \,. \tag{30}$$

The analysis is conducted by EViews 4.0 and by CATS in RATS (see Hansen & Juselius, 1995).

The systems contain (dm^r, y, pd, i_cen), where system 1 (2 and 3) includes dm^{1r} and pd^1 (dm^{3r} and pd^3 as well as dm^{7r} and pd^7, respectively). In addition, a system for M3 is investigated using the variables m3^r, y, R^{bo}, i_cen, and R^{M3}, where R^{bo} is euro area bond yields and R^{M3} is the own rate of simple-sum M3 (see Deutsche Bundesbank, 2001). Augmented Dickey-Fuller and Phillips-Perron tests indicate that all variables in the long-run specification are integrated of order one (see Table 5). To conduct the cointegration analysis in a vector autoregressive (VAR) framework, the lag order of the VAR has to be determined (see Lütkepohl, 1991). Using order selection criteria, the Schwarz criterion (SC) obtains its minimum for order $p = 1$ for systems 1 and 3, whereas the Akaike criterion (AIC) reaches its minimum for $p = 2$ (see Table 6). To be on the safe side $p = 2$ is selected. For system 2 SC criterion estimates a lag order of $p = 1$, whereas the AIC criterion chooses $p = 6$. Nevertheless, $p = 2$ is selected.

The Johansen cointegration trace test is carried out on the assumption that there is an unrestricted intercept in the system. Hence no trend in the cointegrating vector

Table 5: Unit root tests for the variables

Variable	Specification	ADF-t test	PP-test
dm^{1r}	c, t, 1	-3.37*	-2.93
Δ dm^{1r}	c, 1	-3.38**	-4.92***
dm^{3r}	c, t, 1	-2.72	-2.24
Δ dm^{3r}	c	-6.07***	-6.22***
dm^{7r}	c, t, 1, 2	-3.52**	-2.83
Δ dm^{7r}	c, 1	-3.45**	-5.00***
$m3^{r}$	c, t, 1	-2.60	-2.55
Δ $m3^{r}$	c,	-5.26***	-5.81***
pd^{1}	c, t, 2, 4, 5	-2.71	-2.04
Δ pd^{1}	c, 1, 4	-3.57***	-8.07***
pd^{3}	c, t, 1, 5	-2.48	-1.68
Δ pd^{3}	c, 5	-6.37***	-6.96***
pd^{7}	c, t, 1, 4	-2.34	-2.59
Δ pd^{7}	c, 4, 5	-6.96***	-5.81***
y	c, t, 4	-1.85	-2.55
Δ y	c, 1, 2	-3.97***	-7.81***
R^{bo}	c, 1, 2	-1.21	-1.88
Δ R^{bo}	c, 1	-5.31***	-4.60***
R^{M3}	c, 1	-1.95	-1.51
Δ R^{M3}	c	-4.05***	-4.45***
i_cen	c, 1	-1.15	-1.15
Δ i_cen	c	-5.73***	-5.87***

Specification: Select specification of a subset analysis allowing for a maximum lag order of 5. c: intercept term, t: linear trend. ADF-t-test: Augmented Dickey-Fuller t test. PP-test: Phillips-Perron-test using a truncation lag of 3. The information period is 1980Q2 - 2000Q4, except for R^{M3} having the period 1982Q1 - 2000Q4.

is assumed. In addition, the system includes an impulse dummy due to the German unification, which is unity for the second quarter of 1990 and zero elsewhere. The test indicates that there is one cointegrating relationship among the variables for systems 1 and 3 (see Table 7). For systems 2 and 3, the null hypothesis of 1 cointegrating relationship is rejected. Hence two cointegrating vectors are selected.

Residual test statistics indicate that the assumption of the normality of the residuals is not fulfilled for all systems (see Table 8). Since the cointegration theory is asymptotically valid under the i.i.d. assumption of the innovations, this result should not be overvalued. Moreover, there seems to be autocorrelation in the residuals of sys-

Table 6: The lag order of unrestricted VAR is estimated by information criteria

Lag order	1	2	3	4	5	6
	System 1: dm^1r, y, pd^1, i_cen					
AIC-Value	-27.99	-28.20*	-28.16	-28.10	-28.07	-28.14
HQ-Value	-27.70	-27.71*	-27.47	-27.22	-27.00	-26.87
SC-Value	-27.26*	-26.98	-26.45	-25.91	-25.39	-24.97
	Result: $p = 2$					
	System 2: dm^3r, y, pd^3, i_cen					
AIC-Value	-27.37	-27.35	-27.23	-27.29	-27.25	-27.50*
HQ-Value	-27.03*	-26.82	-26.50	-26.36	-26.13	-26.18
SC-Value	-26.51*	-26.01	-25.41	-24.97	-24.45	-24.21
	Result: $p = 2$					
	System 3: dm^7r, y, pd^7, i_cen					
AIC-Value	-26.90	-27.06*	-27.05	-26.90	-26.94	-26.97
HQ-Value	-26.61*	-26.57	-26.37	-26.02	-25.87	-25.70
SC-Value	-26.17*	-25.84	-25.34	-24.71	-24.27	-23.80
	Result: $p = 2$					
	System 4: m3^r, y, R^{bo}, R^{M3}, i_cen					
AIC-Value	-58.58	-58.74	-58.49	-58.38	-58.66	-58.80*
HQ-Value	-58.14*	-57.97	-57.40	-56.97	-56.94	-56.76
SC-Value	-57.46*	-56.81	-55.76	-54.84	-54.33	-53.66
	Result: $p = 2$					

*: Minimum of each criterion. All variables except the interest rate have been transformed into natural logarithms. The information period is 1980Q2 - 2000Q4 for the first 3 systems and 1982Q1 - 2000Q4 for the last system. The unrestricted VAR specification includes an intercept.

tem 3 with unrestricted intercepts, which indicates a misspecification of that system. The problem is solved if the intercept is restricted to lie in the cointegrating space. Under this setting, the cointegrating tests suggest selecting $r = 2$ (see Table 7). The considered autocorrelation tests do not indicate any autocorrelation in the residuals for system 3 with restricted intercepts (see Table 8).

The stability of estimates is more important. It is checked by means of recursive estimation techniques (see Hansen & Johansen, 1999). The null hypothesis is that the cointegration space, which is estimated using the observations up to period t, is identical to full sample estimate. The test statistic asymptotically follows a χ^2 distribution. The starting period is the first quarter of 1990. No problems are apparent for the so-called R-representation, which assumes that the dynamic coefficients are constant and equal

Table 7: Cointegration tests

Null hypothesis	System 1		System 2		System 3				System 4	
	Trace test:	λ_{max} test:	Trace test:	λ_{max} test:	Trace test:	λ_{max} test:	Trace test:	λ_{max} test:	Trace test:	λ_{max} test:
$r = 0$	60.33***	33.90***	66.79***	35.02***	55.42***	28.85**	80.72***	43.46***	84.88***	36.19**
$r = 1$	26.42	21.17*	31.77**	23.23**	26.57	16.17	37.26**	19.07*	48.68**	24.69
$r = 2$	5.25	4.64	8.54	8.40	10.40	8.95	18.19	11.39	23.99	15.56
$r = 3$	0.61	0.61	0.14	0.14	1.45	1.45	6.80	6.80	8.43	8.42
$r = 4$									0.01	0.01

Sample period 1981Q1 - 2000Q4 except system 4, where the period is 1982Q3 - 2000Q4. From the results in the previous paragraph, the order of the VAR was chosen to be 2. The intercept is unrestricted for systems 1, 2, 4, and 3 first block. For system 3 second block the intercept lies in the cointegrating space. *** (**, *): at the 1 % (5 %, 10 %) - level significant. Critical values from table 1 of Osterwald–Lenum (1992).

to the full sample estimate (see Figure 4, Panel a). For the Z-representation, where the dynamic coefficients are re-estimated for each additional observation, instability is indicated for one quarter. In sum, it seems sensible to conclude that no severe instabilities occur for system 1. The stability test does not indicate any instabilities for system 2 (see Figure 4, Panel b). Turning to systems 3 and 4 the stability test for the so-called R-representation give no severe hints of instabilities (see Figure 4, Panels c and d). For the Z-representation, where the dynamic coefficients are re-estimated for each additional observation, instability is indicated for the beginning of the nineties. These effects may capture the influence of German unification and the EMS crisis. Due to the small sample the results should not be overvalued. There seem to be no severe stability problems.

To identify a money demand function in the VAR of systems, as in Coenen and Vega (1999), some restrictions are tested for the loading and cointegrating vector (see Johansen & Juselius, 1992 and Bauwens & Hunter, 2001). The test of the weak exogeneity of variables regarding the long-run relationship restricts the loading vector. It seems sensible that real GDP, price dual and central bank interest rates are weakly exogenous for the cointegrating vector (see Table 9, upper block). The restriction of the exclusion of the central bank interest rate is not rejected. These restrictions are

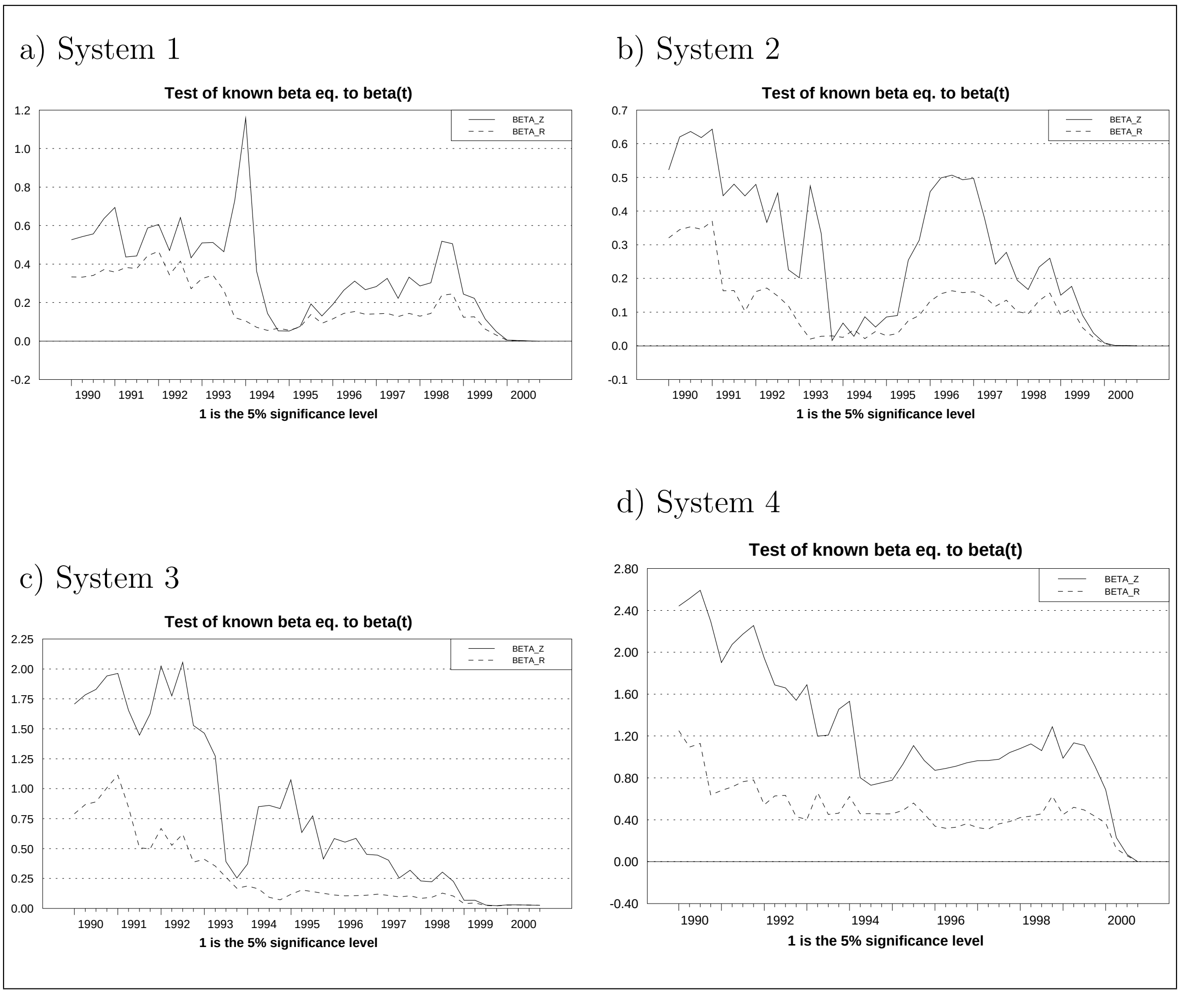

Figure 4: Stability tests of estimated cointegration spaces using the Z-representation and the R-representation for the period 1990Q1 - 2000Q4.

Table 8: Residual test statistics for the systems

Tests	System 1 with DM[1]				System 2 with DM[3]				System 3 with DM[7]				System 4 with M3			
	LM(1)	LM(4)	L-B(20)	Nor.	LM(1)	LM(4)	L-B(20)	Nor.	LM(1)	LM(4)	L-B(20)	Nor.				
	Unrestricted system				Unrestricted system				Unrestr. system, unrestr. interc.				Unrestricted system			
Statistic	17.28	25.91	312.30	61.75	19.15	25.24	322.73	93.11	28.33	17.76	316.44	72.13	39.22	29.33	486.83	26.94
(p-value)	(.37)	(.06)	(.30)	(.00)	(.26)	(.07)	(.14)	(.00)	(.03)	(.34)	(.25)	(.00)	(.04)	(.25)	(.01)	(.00)
	Restricted system				Restricted system				Unrestr. system, restr. interc.				Restricted system			
Statistic	17.97	26.51	319.1	59.40	22.65	32.88	310.9	44.30	23.15	18.85	315.43	68.26	50.33	42.30	539.7	39.23
(p-value)	(.33)	(.05)	(.21)	(.00)	(.12)	(.01)	(.38)	(.00)	(.11)	(.28)	(.21)	(.00)	(.00)	(.02)	(.02)	(.00)
									Restr. system, restr. interc.							
Statistic									22.47	19.22	316.29	70.96				
(p-value)									(.13)	(.26)	(.20)	(.00)				

LM(1) and LM(4) are Lagrange-Multiplier-tests of autocorrelation. L-B(20): Ljung-Box of autocorrelation for 20 autocorrelation matrices. Nor.: Normality test. Tests are described by Hansen & Juselius (1995).

Table 9: Restrictions on loading and cointegrating vectors

Hypo-	System 1: dm^{1r}, y, pd^1, i_cen						Excl.
theses H_i	$B_1 = 0$	$B_2 = 0$	$B_3 = 0$	$B_4 = 0$	$C_4 = 0$	$B_{2,3,4} = 0$ and $C_4 = 0$	$C_4 = 0^{a)}$
Statistic	12.73	0.00	1.86	0.00	0.02	2.19	0.26
p-value	.00	.97	.17	.97	.89	.70	.61
	System 2: dm^{3r}, y, pd^3, i_cen; $r = 2$						Excl.
H_i	$B_{1.} = 0$	$B_{2.} = 0$	$B_{3.} = 0$	$B_{4.} = 0$	$C_{14} = 0$ $B_{12} = B_{21} = 0$ $B_{31} = B_{41} = 0$	$C_{11} = -C_{12}$ $C_{14} = 0$ $B_{12} = B_{21} = 0$ $B_{31} = B_{41} = 0$	$C_{.4} = 0^{b)}$
Statistic	12.47	13.14	3.71	14.87	5.45	16.07	16.36
p-value	.00	.00	.16	.00	.14	.00	.00
	System 3: dm^{7r}, y, pd^7, i_cen; $r = 1$ unrest. interc.						
H_i	$B_1 = 0$	$B_2 = 0$	$B_3 = 0$	$B_4 = 0$			
Statistic	.00	10.77	4.12	6.82			
p-value	.98	.00	.04	.01			
	System 3: dm^{7r}, y, pd^7, i_cen; $r = 2$ restr. interc.						Excl.
H_i	$B_{1.} = 0$	$B_{2.} = 0$	$B_{3.} = 0$	$B_{4.} = 0$	$C_{11} = -C_{12}$ $C_{21} = -C_{22}$ $C_{24} = 0$	$C_{11} = -C_{12}$ $C_{21} = -C_{22}$ $C_{24} = 0$ $B_{11} = B_{22} = 0$ $B_{32} = B_{42} = 0$ $B_{41} = 0$	$C_{.4} = 0^{c)}$
Statistic	9.21	25.24	5.81	4.54	0.33	7.14	18.79
p-value	.01	.00	.06	.10	.57	.14	.00
	System 4: m3^r, y, R^{bo}, i_cen, R^{M3}; $r = 2$ unrestr. interc.						Excl.
H_i	$B_{1.} = 0$	$B_{2.} = 0$	$B_{3.} = 0$	$B_{4.} = 0$	$B_{5.} = 0$	$B_{11} = B_{21} = 0$ $B_{22} = B_{32} = 0$ $B_{52} = 0$ $C_{24} = C_{25} = 0$	$C_{.4} = 0^{d)}$
Statistic	5.09	4.60	2.35	8.59	10.88	6.49	8.69
p-value	.08	.10	.31	.01	.00	.09	.00

The hypotheses are tested by likelihood ratio tests for unrestricted cointegrating vectors (see Johansen and Juselius, 1992, pp. 224-5). The test statistic is asymptotically distributed as $\chi^2(s)$. s number of restrictions. Excl: Exclusion of the i_cen variable from the long-run relationships. a) The restriction test is conducted under the condition that $B_{2,3,4} = 0$. b) The conditions are $C_{11} = -C_{12}$ and $B_{12} = B_{21} = B_{31} = B_{41} = 0$. c) The conditions are $C_{11} = -C_{12}$, $C_{21} = -C_{22}$ and $B_{11} = B_{22} = B_{32} = B_{42} = 0$. d) The conditions are $B_{11} = B_{21} = B_{22} = B_{32} = B_{52} = 0$.

are tested together and the value of the test statistic is 2.19, which has a p-value of 0.70. Thus, the tests indicate that the cointegrating relationship may be interpreted as a long run money demand function. The residual test statistics for the restricted system do not give hints of further problems of the underlying residual assumptions.

When normalised for real Divisia (DM^{1r}), the cointegrating vector takes the following form

$$dm^{1r} = \underset{(25.93)}{1.175}\,y - \underset{(6.95)}{.0603}\,pd^1 \tag{31}$$

where the estimated loading coefficient is

$$\underset{(6.34)}{-.155} \quad \text{in the } \Delta dm^{1r}\text{equation.}$$

All coefficients are statistically significant and have the expected signs.

Identifying restriction tests are repeated for the VEC of system 2. Following Bauwens and Hunter (2001), the identification can be generated by restrictions on the loading vectors. The test results are given in Table 9. The hypothesis of weak exogeneity regarding both cointegration vectors is rejected for all four variables. Selecting the first cointegrating vector as a money demand equation implies some restrictions. The hypothesis is specified in such a way that the i_cen coefficient is zero and the loading coefficients of this cointegrating vector is zero in the price dual, real GDP and i_cen equation. These restrictions are not rejected at the 10 per cent level. If the hypothesis of an income elasticity of unity is additionally tested, the corresponding value of the test statistic is 16.07, which is significant at the 1 per cent level. Therefore, the estimated long-run money demand relationship is

$$dm^{3r} = \underset{(12.53)}{1.36}\,y - \underset{(3.79)}{.11}\,pd^3 \tag{32}$$

and the estimated loading parameter

$$\underset{(4.14)}{-.098} \quad \text{in the } \Delta dm^{3r}\text{equation.}$$

Turning to the VEC of system 3, the hypothesis of weak exogeneity in respect of both cointegrating vectors is rejected at the 5 per cent level for the real money and GDP variable (see Table 9). This hypothesis is not rejected for the interest rate variable at the 10 per cent level. The hypothesis that the money and income coefficients of the cointegrating vectors are equal with different signs is not rejected. If the second cointegrating vector is a money demand equation, the corresponding loading coefficients of the other equations are set at zero. Moreover the loading coefficient of the first cointegrating vector is set to zero in the money equation. These restrictions are not rejected at the 10 per cent test level. It is worth noting that these restrictions identify the system (see Bauwens, Hunter, 2001). On these assumptions, the estimated money demand long-run relationship is:

$$dm^{7r} = 1.00y - \underset{(7.02)}{.063} pd^{7} - \underset{(21.86)}{1.307} \tag{33}$$

and the estimated loading parameter:

$$\underset{(3.97)}{-.106} \quad \text{in the } \Delta dm^{7r} \text{equation.}$$

The coefficients are statistically significant and have the expected signs. In sum, the presented VEC include stable money demand functions.

These results are compared with the evidence for M3. Studies by Coenen and Vega (1999), Brand and Cassola (2000), Calza, Gerdesmeier and Levy (2001) presented evidence of a stable long run M3 money demand function. The studies differ regarding the definition of the opportunity costs of holding money (see Deutsche Bundesbank). The system examined in this study includes two long run relationships, where one is identified as a long-run money demand function (see Table 9, last part). The residual test statistics for the restricted system indicate autocorrelation problems. Nevertheless, to compare the system with the others for the long-run relationship the results are:

$$m3^{r} = \underset{(9.59)}{1.40}\, y - \underset{(3.21)}{1.73}\, R^{bo} \tag{34}$$

and the estimated loading parameter is:

$$\underset{(3.58)}{-.056} \quad \text{in the } \Delta m3^{r} \text{equation.}$$

The long-run relationship confirms the approach of Brand and Cassola (2000) approximating the spread of the bond yields and the own interest rate of M3 by bond yields.

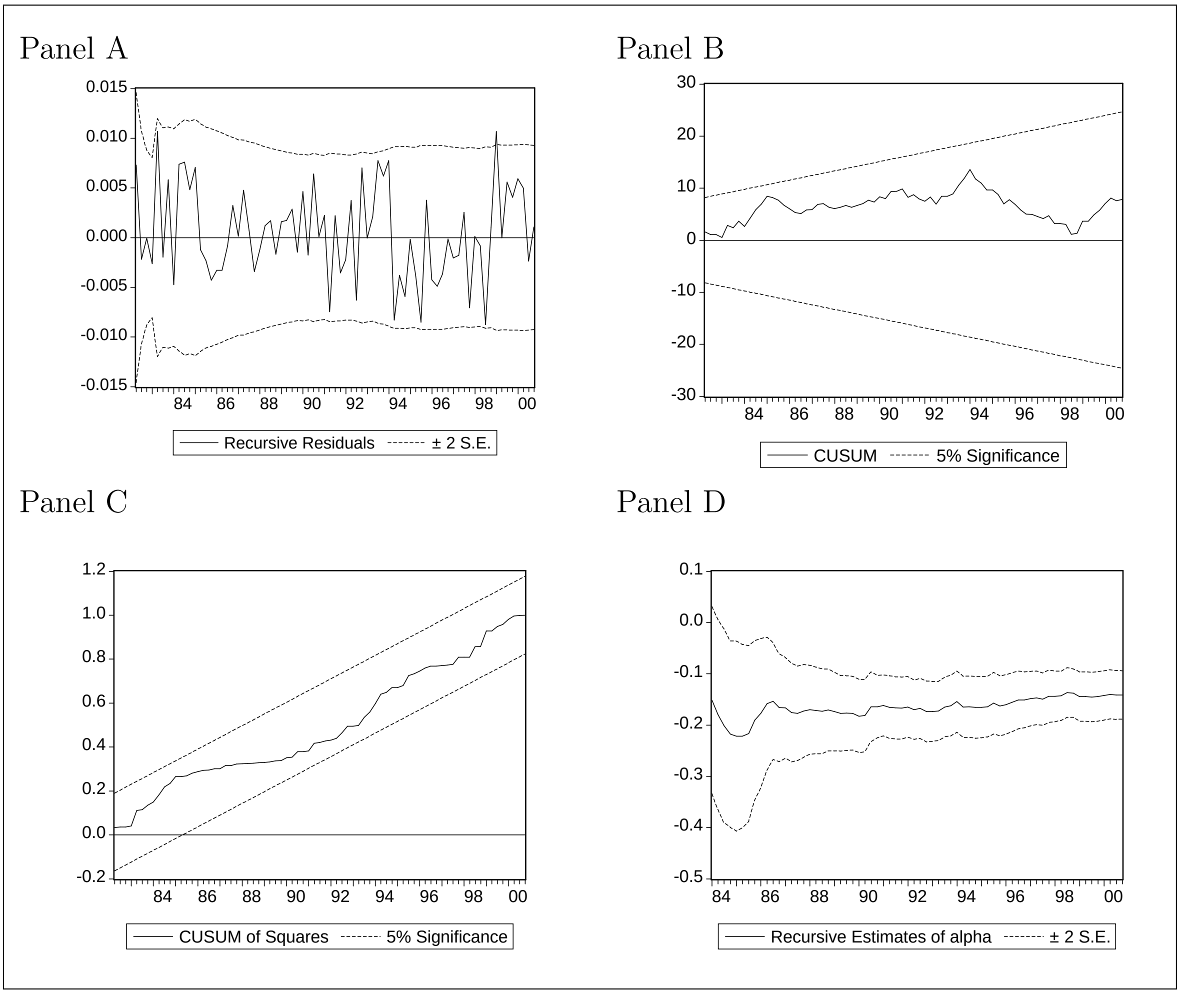

Figure 5: Stability tests for the money demand equation of dm^{1r}; Panel A: Recursive residuals; Panel B: CUSUM-test; Panel C: CUSUMQ-test; Panel D: Recursive estimates of the loading coefficient.

To be in line with the studies by Coenen and Vega (1999) and Brand and Cassola (2000) a single-equation approach is specified, where the dynamic coefficient may be set at zero. Starting with a lag order of two, coefficients which have a small t-value in absolute terms are set stepwise at zero. The preferred specification of system 1 is given in the following equation:

$$\Delta dm_t^{1r} = \underset{(9.70)}{.0083} + \underset{(5.97)}{.339}\,\Delta dm_{t-1}^{1r} - \underset{(1.75)}{.003}\,\Delta i\_cen_{t-1} - \underset{(6.89)}{.141}\,ec_{t-1}^{1} + \underset{(38.9)}{.023}\,Dum903 \quad (35)$$

$$\overline{R^2} \;=\; .606 \quad DW = 2.21 \quad L-B(16) = \underset{(.903)}{9.25} \quad Chow(12) = \underset{(.274)}{1.245}$$

$$J.-B. = \underset{(.781)}{.495} \quad ARCH(1)\underset{(.561)}{.341} \quad RESET(2) = \underset{(.124)}{2.14}$$

$$\text{LMAR(1)} = \underset{(.153)}{2.08} \quad \text{LMAR(1-2)} = \underset{(.319)}{1.16} \quad Heteros. = \underset{(.290)}{1.246}$$

where ec^1 are residuals of the cointegrating relationship (31) and Dum903 is an impulse dummy for German unification.[2] It is unity in 1990Q3 and zero elsewhere. The battery of diagnostic tests does not indicate any problems of the underlying assumptions. The stability tests used do not indicate any instability in this equation (see Figure 5).

The preferred equation of system 2 is:

$$\Delta dm_t^{3r} = \underset{(3.78)}{.282}\,\Delta dm_{t-1}^{3r} - \underset{(1.74)}{.283}\,\Delta y - \underset{(3.94)}{.086}\,ec_{t-1}^{3} + \underset{(32.7)}{.029}\,Dum903 \tag{36}$$

$$\overline{R^2} = .376 \quad DW = 2.13 \quad L-B(16) = \underset{(.967)}{7.84} \quad Chow(12) = \underset{(.810)}{.556}$$

$$J.-B. = \underset{(.000)}{19.5} \quad ARCH(1)\underset{(.430)}{.629} \quad RESET(2) = \underset{(.447)}{.813}$$

$$\text{LMAR(1)} = \underset{(.193)}{1.73} \quad \text{LMAR(1-2)} = \underset{(.428)}{.858} \quad Heteros. = \underset{(.726)}{.634},$$

where ec^3 are residuals of the cointegrating relationship (32). The stability tests used do not indicate any instability in this equation (see Figure 6).

The dynamic money demand function from system 3 is:

$$\Delta dm_t^{7r} = \underset{(4.30)}{.370}\,\Delta dm_{t-1}^{7r} + \underset{(1.72)}{.169}\,\Delta y - \underset{(4.14)}{.100}\,ec_{t-1}^{7} + \underset{(25.8)}{.016}\,Dum903 \tag{37}$$

$$\overline{R^2} = .437 \quad DW = 2.16 \quad L-B(16) = \underset{(.744)}{12.0} \quad Chow(12) = \underset{(.483)}{.974}$$

$$J.-B. = \underset{(.441)}{1.64} \quad ARCH(1)\underset{(.934)}{.007} \quad RESET(2) = \underset{(.825)}{.193}$$

[2] The diagnostic tests are conducted using EViews. L-B(16): Ljung-Box test using 16 autocorrelations. LMAR(·): Lagrange-Multiplier test of autocorrelation using 1 or 1 to 2 autocorrelations. Heteros: Test of heteroskedasticity with cross terms. RESET: Ramsey's non-linearity test. p-values in parentheses.

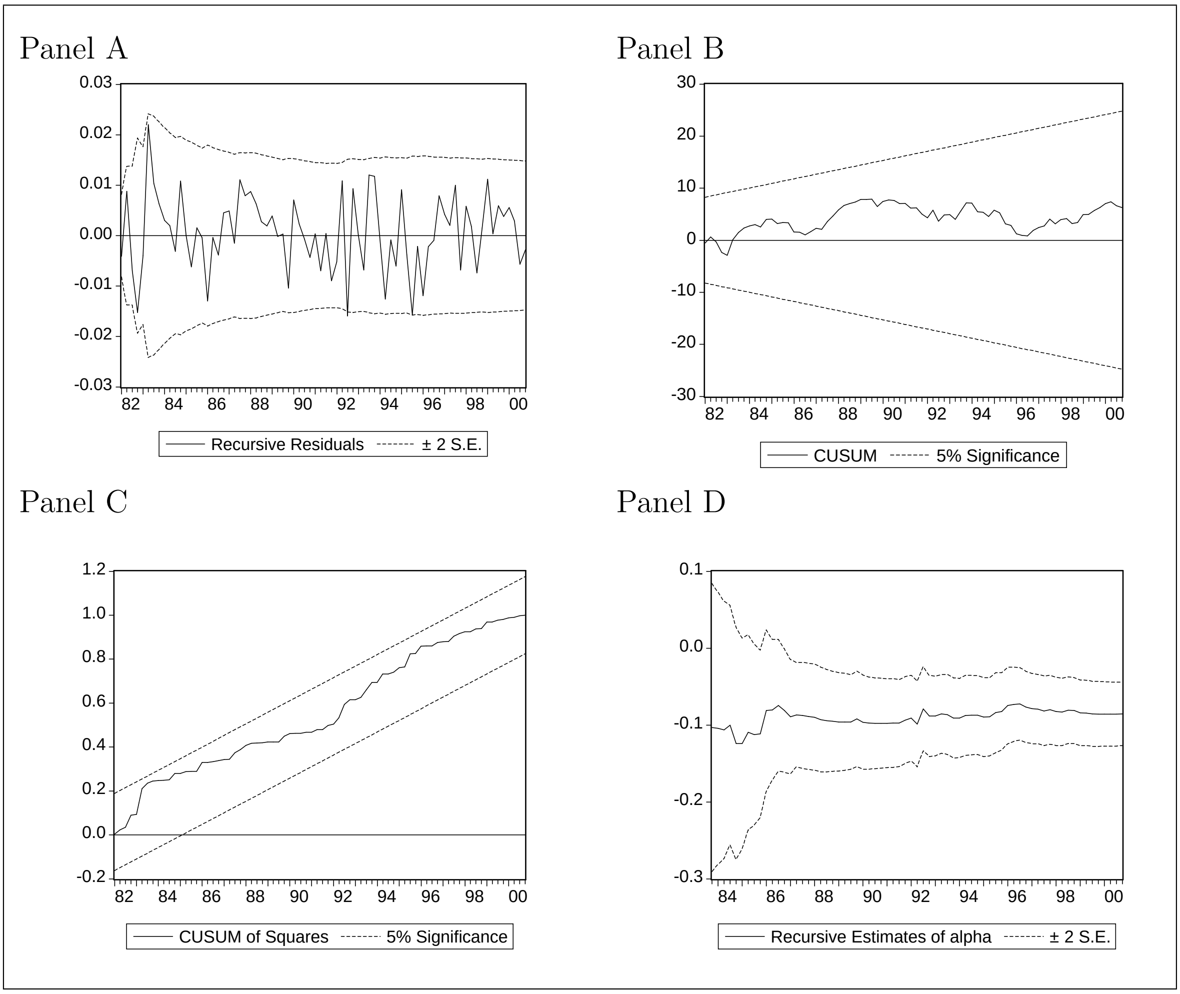

Figure 6: Stability tests for the money demand equation of dm^{3r}; Panel A: Recursive residuals; Panel B: CUSUM-test; Panel C: CUSUMQ-test; Panel D: Recursive estimates of the loading coefficient.

$$\text{LMAR(1)} = \underset{(.111)}{2.60} \quad \text{LMAR(1-2)} = \underset{(.125)}{1.14} \quad Heteros. = \underset{(.373)}{1.099},$$

where ec^7 are residuals of the cointegrating relationship (33). The diagnostic tests considered do not indicate any problems with underlying residual assumptions. The stability tests applied do not indicate any severe instabilities in this equation (see Figure 7).

The single-equation money demand function of system 4 is

$$\Delta m3^r_t = \underset{(3.86)}{.288}\,\Delta m3^r_{t-1} - \underset{(2.07)}{.196}\,\Delta y + \underset{(1.90)}{.471}\,\Delta R^{M3} - \underset{(4.56)}{.051}\,ec^{M3}_{t-1} + \underset{(40.3)}{.020}\,Dum903 \quad (38)$$

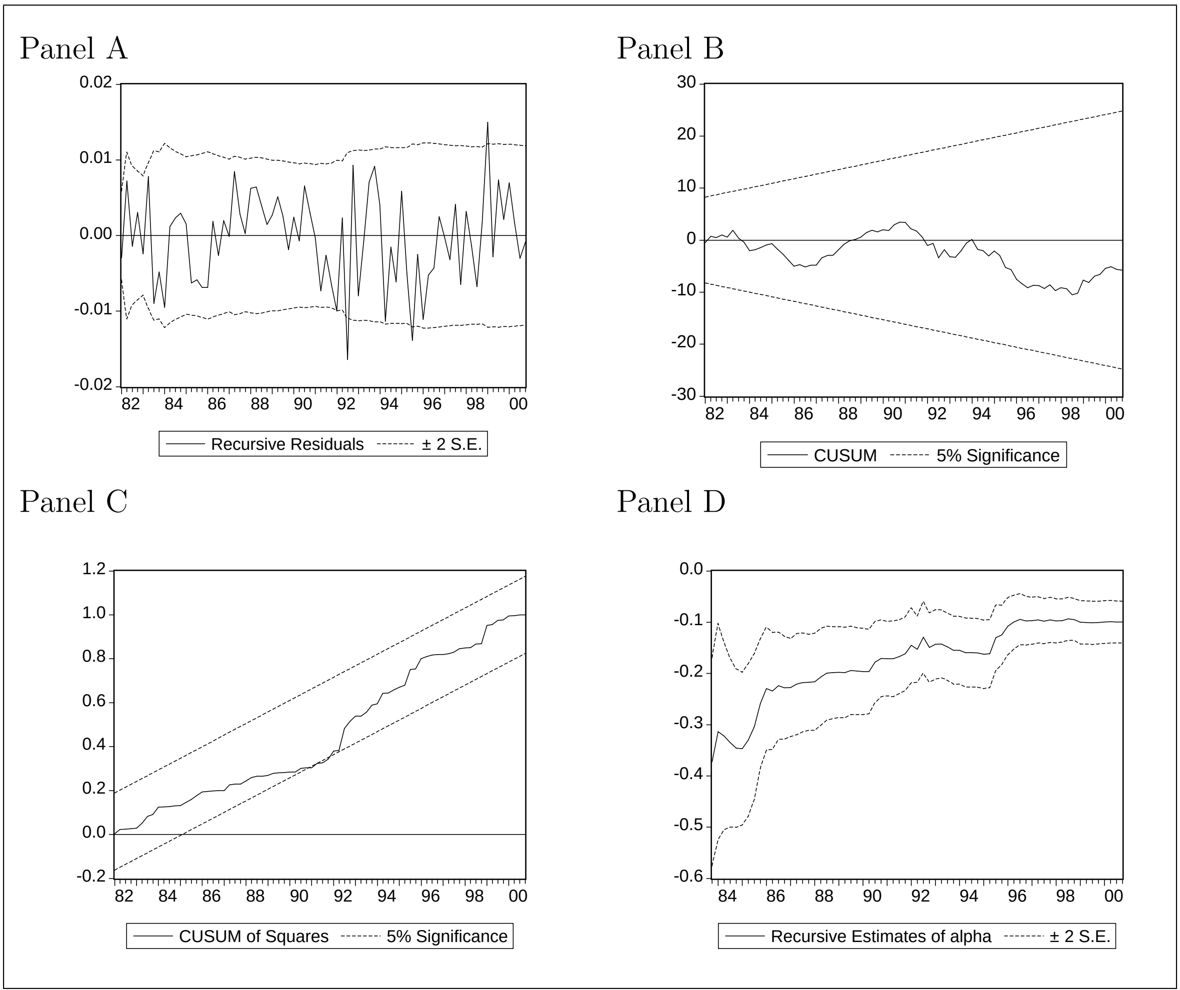

Figure 7: Stability tests for the money demand equation of dm^{7r}; Panel A: Recursive residuals; Panel B: CUSUM-test; Panel C: CUSUMQ-test; Panel D: Recursive estimates of the loading coefficient.

$$\begin{aligned}
\overline{R^2} &= .470 \quad DW = 2.05 \quad L-B(16) = \underset{(.562)}{14.5} \quad Chow(12) = \underset{(.259)}{1.33} \\
J.-B. &= \underset{(.426)}{1.71} \quad ARCH(1)\underset{(.476)}{.513} \quad RESET(2) = \underset{(.070)}{2.77} \\
\text{LMAR(1)} &= \underset{(.735)}{.115} \quad \text{LMAR(1-2)} = \underset{(.905)}{.100} \quad Heteros. = \underset{(.168)}{1.50},
\end{aligned}$$

where ec^{M3} are residuals of the cointegrating relationship (34). The diagnostic tests applied do not suggest that any problems are posed by underlying residual assumptions. The stability tests considered do not indicate any severe instabilities in this equation (see Figure 8). In sum, the estimated single-equations are stable.

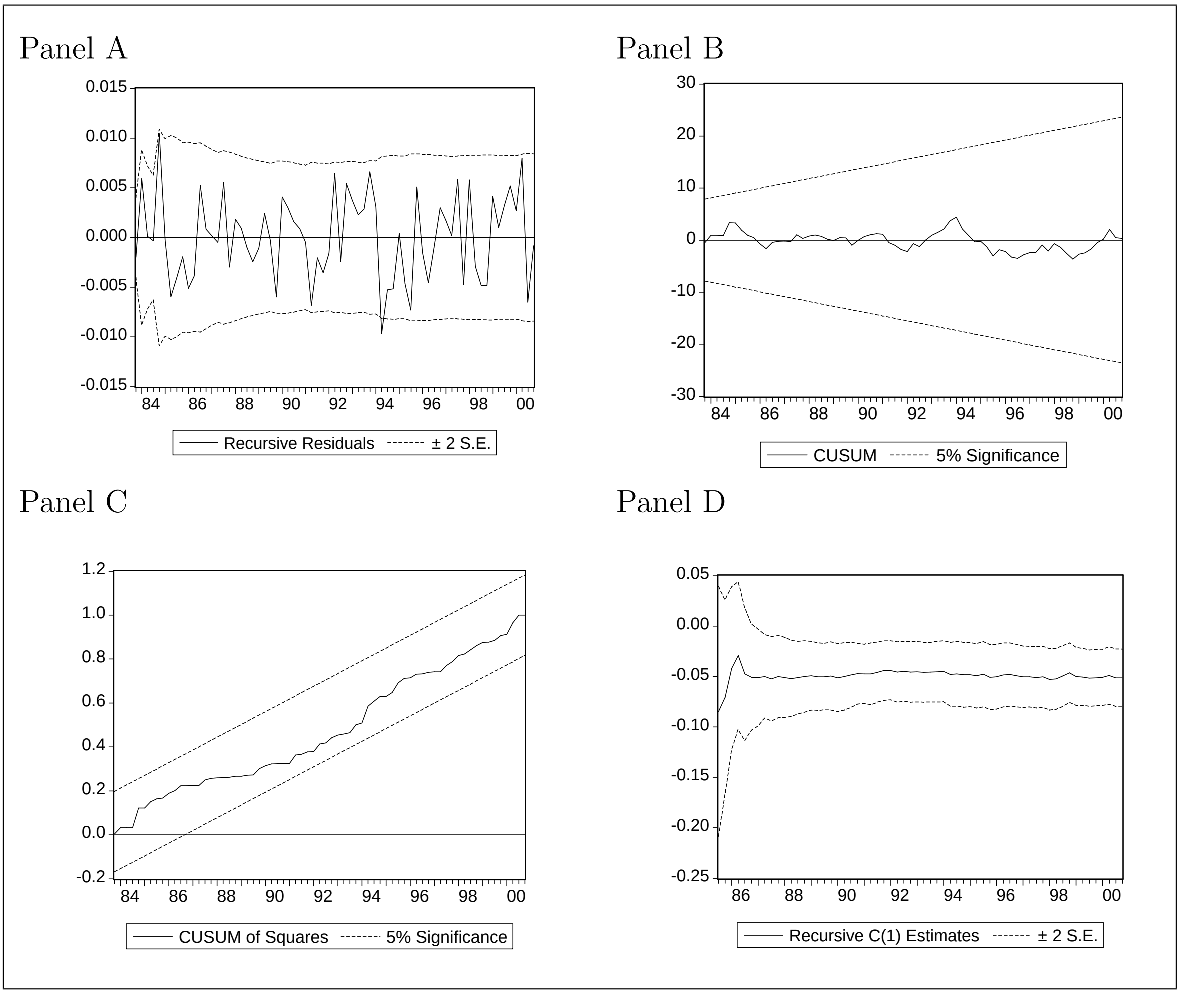

Figure 8: Stability tests for the money demand equation of $m3^r$; Panel A: Recursive residuals; Panel B: CUSUM-test; Panel C: CUSUMQ-test; Panel D: Recursive estimates of the loading coefficient.

In line with Stracca (2001a) for the DM^2 aggregate, and with Coenen & Vega (1999) and Brand & Cassola (2000) for the M3 aggregate, the long-run income elasticity of the real money function is greater than unity. Only for DM^7 is this elasticity unity. The opportunity cost variables of money holding are different from the M3 money demand equations. Coenen & Vega (1999) include the spread between the long-run and short-term interest rates and the inflation rate, whereas Brand & Cassola (2000) estimate a relationship with long-run and short-term interest rates. Moreover, the estimated long-run demand functions of Calza, Gerdesmeier and Levy (2001) contain the spread between the short run interest rate and a calculated own interest rate of the

M3 aggregate, whereas the Deutsche Bundesbank (2001) presents an equation including the spread between the long-run interest rate and the own rate of M3. The result of our specification search, the long-run interest rate seems to be sufficient. These different specifications show the difficulties presented by finding the right measure. For the Divisia aggregates, the coefficient of the price dual variables has the expected sign. In contrast to Stracca (2001a), the empirical results give no hint of additionally including the squared price dual variable. The loading coefficient of the long-run relationships is negative. Its values are in line with estimates of M3 money demand functions.

Controllability is tested directly by the significance of the i_cen variable in the money demand relation. Evidence is presented that the variable can be excluded. It is worth noting that the single-equation results presented have in common that the central bank interest rate i_cen is not included in the long-run equation. In the dynamic part, the demand equation of DM^1 contains this interest rate, however not in the equations for DM^3 and DM^7. On the other hand, the exclusion of the i_cen variable is tested for all relationships. The hypothesis checks the necessary condition of controllability that instrument and target are cointegrated. The test is conducted under identification restrictions. In the case of two cointegrating relationships, the exclusion restriction is rejected (see Table 9, right part). For the last three systems, this evidence indicates that the central bank may indirectly influence the money growth rate in the desired direction by changing central bank interest rates.

Adopting the approach that unexpected shocks of the central bank interest rates are the variable to affect money growth, controllability may be tested by impact matrices (30), where the system results are used. The estimated impact matrices are presented in Table 10. It is apparent that central bank interest rate shocks are negative in the real money equation of system 1, as theoretically expected for controllability. For system 2, a significantly negative effect is found, whereas the influence of the corresponding shocks in the other systems seems to be insignificant. These results indicate that the ECB has only limited potential for controlling these monetary aggregates. However, in contrast to the evidence of Johansen and Juselius (2001) for the US economy, the signs of the shocks are as expected for the Divisia aggregates.

Table 10: Estimates of impact matrices of the shocks

Equation	System 1					System 2				
	dm^{1r}	y	pd^1	i_cen		dm^{3r}	y	pd^3	i_cen	
dm^{1r}	.000	1.67	−.071	−.004	dm^{3r}	.006	1.11	−.091	−.024	
	(.000)	(4.00)	(4.00)	(1.05)		(.147)	(2.04)	(3.30)	(2.04)	
y	.000	1.45	.000	−.001	y	−.398	1.45	.007	−.017	
	(.000)	(4.58)	(.029)	(.226)		(1.36)	(3.78)	(.339)	(2.13)	
pd^1	.000	.544	1.18	.057	pd^3	−4.02	4.77	.728	.019	
	(.000)	(.109)	(5.54)	(1.16)		(1.31)	(1.19)	(3.59)	(.221)	
i_cen	.000	16.8	.791	1.65	i_cen	−10.7	8.22	2.24	.140	
	(.000)	(.535)	(.593)	(5.35)		(1.18)	(.691)	(3.73)	(.549)	
	System 3					System 4				
	dm^{7r}	y	pd^7	i_cen		$\mathrm{m3}^r$	y	R^{bo}	i_cen	R^{M3}
dm^{7r}	.729	2.67	.259	−.214	$\mathrm{m3}^r$	.473	.836	−1.45	.018	−.989
	(.208)	(.682)	(.658)	(.816)		(1.87)	(3.28)	(2.73)	(.057)	(.891)
y	.709	2.59	.252	−.158	y	.087	1.16	−.013	−.155	.051
	(.265)	(.868)	(.838)	(.788)		(.382)	(5.07)	(.027)	(.537)	(.051)
pd^7	−.323	−1.18	−.115	.906	R^{bo}	−.159	.313	.635	−.099	.469
	(.022)	(.073)	(.070)	(.832)		(1.38)	(2.69)	(2.63)	(.680)	(.927)
i_cen	−.944	−3.45	−.336	2.65	i_cen	−.543	.715	−.219	1.15	−.591
	(.022)	(.073)	(.070)	(.832)		(2.22)	(2.91)	(.428)	(3.73)	(.553)
					R^{M3}	−.235	.382	.260	.276	.070
						(1.99)	(3.21)	(1.05)	(1.85)	(.135)

The impact matrices are calculated for system 1 using restrictions on the cointegrating and loading vector and for systems 2, 3 and 4 only using restrictions on the cointegrating vectors. The estimated t-values are in parentheses.

5 The Importance of Liquidity

Money has a role to play as an information variable for monetary policy. To ascertain whether money contains any marginal information about future realisations of variables which monetary policy-makers care about, two approaches are investigated: on the one hand, liquidity for the IS-curve and, on the other hand, liquidity in an inflation equation.

5.1 The IS-curve approach

Theoretical questions concerning the direct money channel of the monetary transmission process are raised by Nelson (2001). He presents an IS equation for log output y_t

$$y_t = -c_1 r_t + E_t y_{t+1}, \qquad (39)$$

where r_t is the real interest rate, which is in some cases approximated by a short-term real interest rate (r_t^s). In that case, iterations on the IS function produce:

$$\begin{aligned} y_t &= -c_1 r_t^s + E_t y_{t+1} \\ &= -c_1 r_t^s - c_1 E_t r_{t+1}^s + E_t y_{t+2} \\ &= \ldots \\ &= -c_1 r_t^l, \end{aligned} \tag{40}$$

where $r_t^l = E_t \sum_{j=0}^{\infty} r_{t+j}^s$ is a long-run real interest rate, according to the expectations theory of the term structure. The last relationship stresses that, for the forward looking IS equation, the long-run real interest rate matters (see Rotemberg and Woodford 1997, 1999).

Noting that money demand depends not only on a short-term interest rate, but also on a range of interest rates (see Friedman, 1956) it may be specified as a semi-logarithmic long-run money demand function and a partial-adjustment formulation of dynamic adjustment

$$m_t - p_t = c_2 y_t - c_3 R_t^l + c_4(m_{t-1} - p_{t-1}), \tag{41}$$

where lower cases denote logs $c_2 > 0$, $c_3 > 0$, $0 \leq c_4 < 1$ and $R_t^l = E_t \sum_{j=0}^{\infty}(\Delta p_{t+j+1} + r_{t+j}^s)$ is the nominal long-run rate. Assuming $c_4 \approx 1$ and using $y_t = -c_1 r_t^l$, the money demand function reads as:

$$\Delta(m-p)_t \approx -c_1 c_2 r_t^l - c_3 R_t^l \tag{42}$$

The change in real money depends negatively on both the real and the nominal long-run interest rate. If inflation persistence makes r_t^l and R_t^l highly correlated, the $\Delta(m-p)_t$ will be a good indicator of the real long-term yield r_t^l, which is the crucial interest rate for aggregate demand. Moreover, Nelson (2001) presents a general equilibrium model to strengthen his position. Quoting the work of Rudebusch and Svensson (1999, 2000) he suggests the simplified backward-looking IS-equation:

$$y_t = c_0 + c_1 y_{t-1} + c_2 r_t + c_3 \Delta(m-p)_{t-1} \tag{43}$$

The last term will be statistically significant, if the prior change in real balances contains information about the next period's output not yet present in lagged output and current short-term real interest rates. For the UK and US economies, he finds evidence in favour of a significant effect of real money changes.

The analysis of the information content of money can be carried out using the IS-curve approach of Rudebusch and Svensson (2000) and Nelson (2001). Rudebusch and Svensson (2000) have recently argued that M2 does not enter significantly into an estimated IS-curve for the US economy. The estimated model is as follows:

$$ygap_t = \delta_0 + \delta_1 ygap_{t-1} + \delta_2 r^{real}_{t-1} + \delta_3(L)\Delta(m-p)_{t-1} + u_t,$$

where $ygap$ is the output gap, r^{real} is a sum of lags of interest rates minus the inflation rate, e.g. $r^{real}_t = \sum_{j=0}^{3} R_{t-j}/400 - (p_t - p_{t-4})$, where R is either a money market interest rate or a bond yield rate. Furthermore, Stracca (2001a) proposes using "excess liquidity". This indicator appears to be of interest for analysis because several interest rates enter into its determination, as well as opportunity cost. The estimated model is:

$$ygap_t = \delta_0 + \delta_1 ygap_{t-1} + \delta_2 r^{real}_t + \delta_3 exliq_{t-1} + u_t,$$

where $exliq$ is an excess liquidity indicator based on the disequilibrium of the money demand market. It is approximated by the residuals of the estimated long-run money demand function for the Divisia aggregates (31), (32) and (33), or for M3

$$m3^r_t = 1.30y_t - 1.76(R^{bo}_t - R^{M3}_t),$$

where R^{bo} government bond yields and R^{m3} own rate of simple-sum M3 (see Deutsche Bundesbank, 2001). Potential output is estimated via a Hodrick-Prescott (HP) filter and expanded exponential smoothing[3] (see Tödter, 2000a, 2000b). Figure 9 exhibits the development of the series and the implied output gaps. The peaks and troughs of the gaps are more or less in the same quarter. However, the output gap of expanded exponential smoothing is more volatile than the series constructed by the HP filter.

To start the empirical analysis, the IS-curve is estimated without any money variable (see Table 11). The description of the dynamics of the equation needs, in one case, the lagged output gap of order 5, elsewhere the lagged output gap of order 1 is sufficient. The estimated equations are free of autocorrelation. The hypothesis of

[3] A brief review of the methods is given in the Appendix.

Table 11: Estimates of the IS-curve using the output gap as endogenous variable

Variable	HP filter				EES filter			
c	.001	.001	.001	.001	.003	.002	.002	.002
	(1.24)	(1.23)	(1.05)	(1.50)	(2.30)	(2.51)	(2.36)	(2.41)
ygap_{t-1}	.836	.842	.835	.842	.923	1.01	.926	.935
	(12.2)	(12.3)	(13.1)	(12.2)	(18.0)	(20.3)	(18.1)	(18.7)
ygap_{t-5}						−.121		
						(2.18)		
$\sum R^{bo} - \Delta_4 pc$	−.010				−.019			
	(1.46)				(2.62)			
$\sum R^{bo} - \Delta_4 pb$		−.010				−.014		
		(1.38)				(1.71)		
$\sum R^{mo} - \Delta_4 pc$			−.010				−.019	
			(1.21)				(2.61)	
$\sum R^{mo} - \Delta_4 pb$				−.011				−.020
				(1.55)				(2.54)
$\overline{R}^2$	.694	.693	.695	.695	.870	.875	.870	.870
DW	1.87	1.87	1.87	1.88	1.83	2.06	1.85	1.85
LMAR(1-2)	.221	.134	.196	.175	.287	.498	.232	.218
	(.802)	(.875)	(.822)	(.840)	(.752)	(.610)	(.793)	(.805)
J.-B.	6.81	6.74	6.28	6.17	3.10	4.52	2.69	2.45
	(.033)	(.034)	(.043)	(.046)	(.212)	(.104)	(.260)	(.294)

ygap: Difference between log GDP and trend of log GDP, as estimated by Hodrick-Prescott (HP)-filter or expanded exponential smoothing (EES) filter. $\sum R^{bo} - \Delta_4 pb$: Real interest rate variable that is $\sum_{j=1}^{4} R^{bo}_{t-j}/400 - \Delta_4 pb_{t-1}$ where R^{bo} is the euro area bond yields and R^{mo} the euro area money market rate. pb GDP deflator. pc Harmonised index of consumer prices. Heteroskedasticity consistent covariance estimated t-values in parentheses. LMAR(1-2): Lagrange-multiplier test of autocorrelation of 1 and 2 lags. J.-B.: Jarque-Bera-test of normality. Diagnostic statistics have p-value in parentheses. Estimation period: 1982Q1-2000Q4.

normal distributed residuals is often rejected at the 5 per cent significance level. The coefficients of the different real interest rate variables are all negative as expected. If the EES filter is used, they are significantly negative.

The results change if the annual growth rate of money is analysed. At first, the output gap determined by the Hodrick-Prescott filter is examined (see Table 12). Owing to delays, the lagged money changes are specified and their coefficients are significant regardless of which money concept is used. The signs of the coefficients are positive, as expected. However, the coefficients of the real interest rate variables are positive and significant for the equations including DM^1 and DM^7, which is in contrast to the theory presented. The residuals does not seem to be normally distributed, which complicate

Table 12: Estimates of the IS-curve, using the output gap as endogenous variable constructed by the Hodrick-Prescott-filter, additionally explained by real money change

Variable	Simple sum M3				Divisia DM[1]				Divisia DM[3]				Divisia DM[7]			
c	−.003	−.003	−.002	−.002	−.006	.006	−.004	−.004	−.005	−.006	−.004	−.004	−.006	−.000	−.005	.005
	(1.25)	(1.54)	(1.08)	(1.24)	(2.37)	(2.94)	(2.03)	(2.77)	(2.90)	(3.18)	(2.83)	(2.91)	(2.61)	(3.16)	(2.44)	(3.03)
ygap_{t-1}	.790	.785	.793	.792	.774	.758	.782	.773	.815	.799	.817	.808	.811	.795	.813	.800
	(12.0)	(10.6)	(12.1)	(10.8)	(12.3)	(13.3)	(12.4)	(12.9)	(11.8)	(11.5)	(11.8)	(11.5)	(13.7)	(16.8)	(13.7)	(17.0)
$\sum R^{bo} - \Delta_4 pc$	.003				.020				.016				.021			
	(.252)				(1.63)				(2.21)				(1.83)			
$\sum R^{bo} - \Delta_4 pb$		.010				.025				.024				.027		
		(.483)				(2.31)				(2.58)				(2.58)		
$\sum R^{mo} - \Delta_4 pc$			−.001				.013				.011				.018	
			(.089)				(1.18)				(1.63)				(1.60)	
$\sum R^{mo} - \Delta_4 pb$				−.000				.015				.015				.021
				(.017)				(1.66)				(1.80)				(2.31)
$\Delta_4 mr_{t-2}$	.074	.076	.068	.069	.103	.107	.091	.092	.079	.088	.072	.077	.104	.109	.099	.103
	(2.21)	(2.29)	(2.12)	(2.10)	(3.36)	(4.51)	(3.09)	(4.40)	(4.54)	(4.59)	(4.29)	(4.17)	(3.84)	(4.48)	(3.66)	(4.32)
$\overline{R}^2$	.709	.710	.709	.709	.732	.734	.727	.728	.725	.729	.722	.724	.742	.745	.739	.741
DW	1.91	1.91	1.91	1.91	2.02	2.00	2.00	1.99	2.03	2.02	2.01	2.00	2.17	2.15	2.15	2.13
LMAR(1-2)	.129	.134	.126	.126	.525	.411	.359	.290	.429	.424	.308	.285	2.13	1.87	1.80	1.57
	(.879)	(.875)	(.882)	(.882)	(.594)	(.665)	(.700)	(.749)	(.653)	(.656)	(.736)	(.753)	(.126)	(.162)	(.173)	(.215)
J.-B.	6.60	6.78	6.29	6.38	10.4	11.2	10.1	10.6	3.16	3.40	3.31	3.55	18.4	20.6	17.1	18.6
	(.037)	(.034)	(.043)	(.041)	(.006)	(.004)	(.006)	(.005)	(.206)	(.182)	(.191)	(.169)	(.000)	(.000)	(.000)	(.000)

ygap: Difference between log GDP and trend of log GDP, estimated by Hodrick-Prescott-filter. $\sum R^{bo} - \Delta_4 pb$: Real interest rate variable that is $\sum_{j=1}^{4} R^{bo}_{t-j}/400 - \Delta_4 pb_{t-1}$ where R^{bo} is the euro area bond yields and R^{mo} the euro area money market rate. pb GDP deflator. pc Harmonised index of consumer prices. $\Delta_4 mr$ annual change in real Divisia M3 or real simple-sum M3 using GDP deflator. Heteroskedasticity consistent covariance estimated t-values in parentheses. LMAR(1-2): Lagrange-multiplier test of autocorrelation of 1 and 2 lags. J.-B.: Jarque-Bera-test of normality. Diagnostic statistics have p-value in parentheses. Estimation period: 1982Q1-2000Q4.

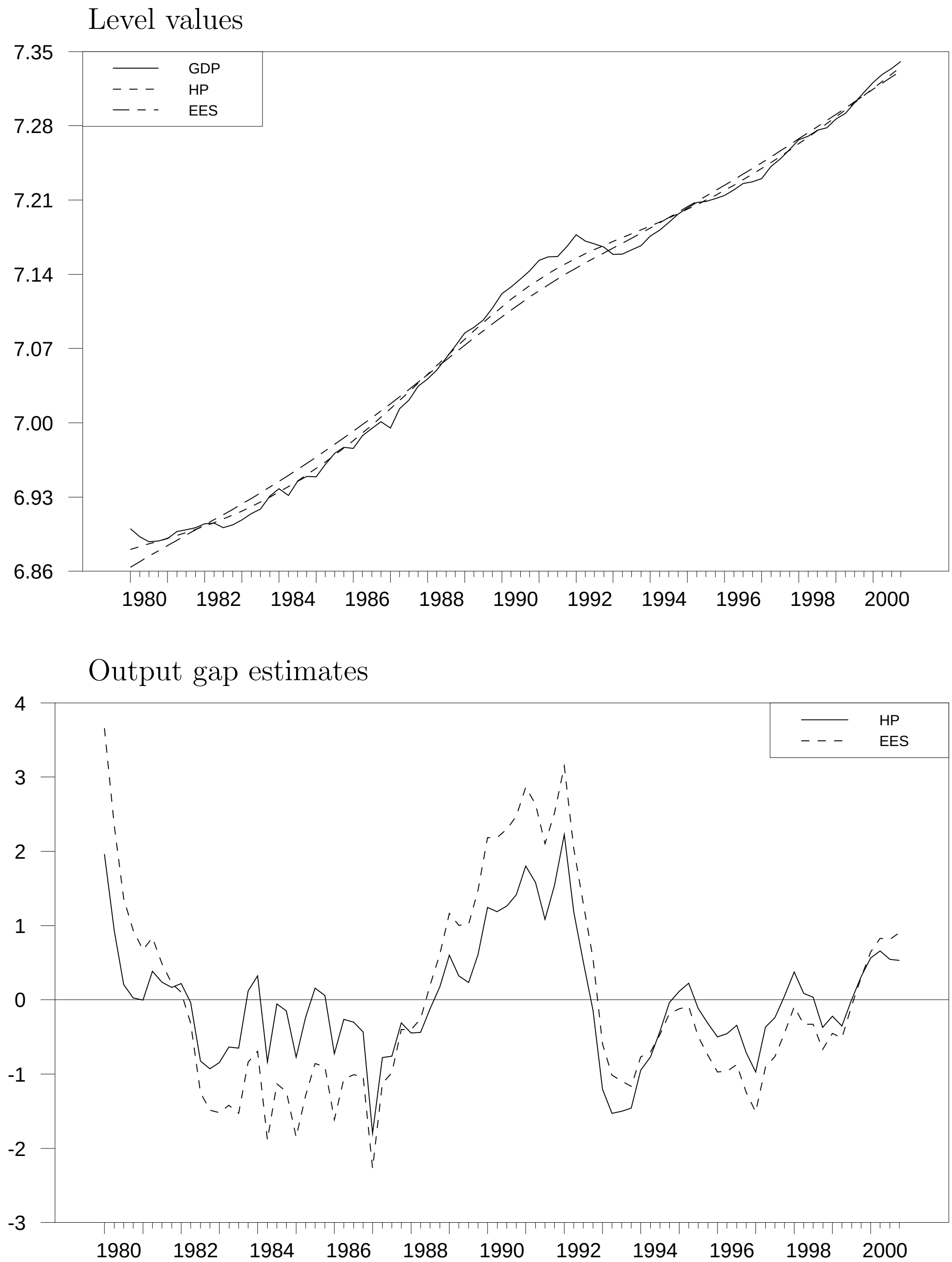

Figure 9: Levels of the log real GDP and its trend estimates HP (Hodrick-Prescott filter) and EES (extended exponential smoothing) (upper panel); Gap between real GDP and its trend estimates in per cent, 1980-2000 (lower panel).

Table 12 continued: Estimates of the IS-curve, using the output gap as endogenous variable constructed by the expanded exponential smoothing filter, additionally explained by real money change

Variable	Simple sum M3				Divisia DM[1]				Divisia DM[3]				Divisia DM[7]			
c	−.002	−.004	−.003	−.003	−.003	−.006	−.007	−.004	−.004	−.005	−.003	−.004	−.006	−.007	−.006	−.006
	(1.72)	(1.86)	(1.58)	(1.66)	(1.66)	(3.13)	(3.17)	(3.36)	(1.95)	(2.41)	(2.03)	(2.50)	(3.19)	(3.43)	(3.35)	(3.45)
ygap_{t-1}	.859	.857	.861	.861	.857	.842	.869	.848	.896	.887	.896	.890	.910	.900	.907	.898
	(18.8)	(18.5)	(18.8)	(18.5)	(22.9)	(22.7)	(22.6)	(22.7)	(18.5)	(18.2)	(18.4)	(18.2)	(29.1)	(29.8)	(28.6)	(29.3)
$\sum R^{bo} - \Delta_4 pc$	−.004				.017				.007				.017			
	(.517)				(1.99)				(.926)				(2.05)			
$\sum R^{bo} - \Delta_4 pb$		.000				.025				.016				.024		
		(.002)				(2.26)				(1.61)				(2.37)		
$\sum R^{mo} - \Delta_4 pc$			−.006				.011				.004				.014	
			(.840)				(1.65)				(.587)				(2.04)	
$\sum R^{mo} - \Delta_4 pb$				−.004				.017				.010				.020
				(.459)				(1.93)				(1.26)				(2.28)
$\Delta_4 mr_{t-2}$	.104	.111	.100	.104	.131	.142	.121	.129	.085	.097	.080	.089	.123	.130	.120	.127
	(3.31)	(3.27)	(3.18)	(3.10)	(5.21)	(5.07)	(5.56)	(5.40)	(4.01)	(4.19)	(4.10)	(4.24)	(5.26)	(5.52)	(5.33)	(5.42)
$\overline{R}^2$	.880	.881	.881	.881	.892	.894	.891	.892	.883	.885	.883	.884	.897	.898	.896	.897
DW	1.92	1.92	1.93	1.93	2.10	2.11	2.08	2.08	1.98	1.99	1.98	1.98	2.29	2.30	2.27	2.28
LMAR(1-2)	.080	.085	.076	.078	.801	.908	.602	.627	.085	.114	.071	.078	3.44	3.66	3.03	3.15
	(.923)	(.918)	(.927)	(.925)	(.453)	(.408)	(.551)	(.537)	(.919)	(.892)	(.932)	(.925)	(.037)	(.031)	(.055)	(.049)
J.-B.	2.73	2.91	2.43	2.57	5.12	6.45	4.89	5.84	1.00	1.32	.945	1.15	8.84	11.2	8.54	10.5
	(.255)	(.233)	(.296)	(.277)	(.077)	(.040)	(.087)	(.054)	(.606)	(.518)	(.624)	(.564)	(.012)	(.004)	(.014)	(.005)

ygap: Difference between log GDP and trend of log GDP, estimated by expanded exponential smoothing. Furthermore, see Table 12.

Table 13: Estimates of the IS-curve, using the output gap as endogenous variable, constructed by the Hodrick-Prescott-filter, additionally explained by money demand disequilibrium (money overhang)

Variable	Simple sum M3				Divisia DM[1]				Divisia DM[3]				Divisia DM[7]			
c	.001	.001	.001	.001	.002	.002	.001	.001	.087	.076	.085	.073	.006	.006	.006	.006
	(.546)	(.352)	(.656)	(.490)	(1.24)	(1.12)	(1.24)	(1.16)	(3.58)	(3.22)	(3.62)	(3.26)	(3.14)	(3.04)	(3.46)	(3.38)
ygap_{t-1}	.878	.878	.877	.879	.861	.871	.862	.873	.859	.881	.855	.876	.810	.827	.805	.823
	(12.5)	(12.4)	(12.5)	(12.5)	(13.5)	(13.3)	(13.6)	(13.5)	(14.3)	(14.2)	(14.3)	(14.3)	(13.3)	(13.6)	(13.4)	(13.7)
$\sum R^{bo} - \Delta_4 pc$	−.006				−.026				−.046				−.027			
	(.599)				(2.19)				(3.45)				(2.67)			
$\sum R^{bo} - \Delta_4 pb$		−.036				−.024				−.044				−.029		
		(.280)				(1.91)				(3.05)				(2.57)		
$\sum R^{mo} - \Delta_4 pc$			−.006				−.028				−.043				−.029	
			(.599)				(2.47)				(3.54)				(3.00)	
$\sum R^{mo} - \Delta_4 pb$				−.005				−.026				−.041				−.031
				(.421)				(2.19)				(3.15)				(2.92)
coin-m	.056	.055	.055	.054	.067	.056	.075	.061	.084	.072	.081	.070	.071	.070	.077	.076
	(1.67)	(1.65)	(1.67)	(1.63)	(2.07)	(1.82)	(2.29)	(2.02)	(3.54)	(3.17)	(3.57)	(3.21)	(3.27)	(3.20)	(3.51)	(3.45)
$\overline{R}^2$	.700	.699	.700	.699	.707	.703	.712	.707	.736	.727	.737	.729	.730	.728	.736	.734
DW	1.87	1.87	1.87	1.88	1.86	1.88	1.87	1.89	2.06	2.06	2.07	2.07	1.90	1.93	1.93	1.96
LMAR(1-2)	.203	.197	.190	.181	.295	.201	.211	.139	.433	.452	.561	.554	.153	.146	.179	.238
	(.817)	(.822)	(.828)	(.835)	(.745)	(.818)	(.810)	(.871)	(.650)	(.638)	(.573)	(.577)	(.859)	(.864)	(.837)	(.789)
J.-B.	3.50	3.96	3.14	3.55	4.81	4.95	3.48	3.75	3.69	3.93	3.76	4.13	6.08	5.80	4.87	4.65
	(.174)	(.138)	(.208)	(.169)	(.090)	(.084)	(.176)	(.154)	(.158)	(.140)	(.153)	(.127)	(.048)	(.055)	(.087)	(.098)

ygap: Difference between log GDP and trend of log GDP, estimated the by Hodrick-Prescott-filter. coin-m: Residuals of the estimated long-run money demand function using Divisia M3 or simple-sum M3. Estimation period: 1982Q1-2000Q4 for equations using Divisia M3 and 1983Q2-2000Q4 for equations using simple-sum M3. Further comments see Table 12.

Table 13 continued: Estimates of the IS-curve, using the output gap as endogenous variable constructed by the expanded exponential smoothing filter, additionally explained by money demand disequilibrium (money overhang)

Variable	Simple sum M3				Divisia DM[1]				Divisia DM[3]				Divisia DM[7]			
c	.001	.001	.001	.001	.002	.002	.001	.001	.090	.074	.086	.071	.008	.007	.008	.007
	(1.06)	(.835)	(.832)	(.596)	(1.81)	(1.66)	(1.47)	(1.34)	(5.02)	(4.33)	(4.49)	(3.87)	(4.42)	(3.91)	(4.28)	(3.76)
ygap_{t-1}	1.48	1.04	1.04	1.04	1.03	1.05	1.03	1.04	.981	1.01	.973	1.00	.935	.962	.923	.951
	(19.8)	(20.8)	(19.5)	(20.5)	(20.0)	(21.0)	(19.6)	(20.5)	(17.6)	(19.5)	(16.9)	(18.8)	(18.2)	(19.5)	(17.3)	(18.5)
ygap_{t-5}	−.133	−.135	−.130	−.133	−.157	−.159	−.141	−.144	−.073	−.091	−.050	−.070	−.088	−.098	−.065	−.076
	(2.10)	(2.13)	(2.02)	(2.07)	(2.59)	(2.64)	(2.37)	(2.43)	(1.18)	(1.58)	(.812)	(1.23)	(1.49)	(1.71)	(1.11)	(1.34)
$\sum R^{bo} - \Delta_4 pc$	−.009				−.033				−.053				−.037			
	(1.03)				(3.37)				(4.79)				(4.11)			
$\sum R^{bo} - \Delta_4 pb$		−.007				−.030				−.048				−.037		
		(.754)				(2.99)				(4.20)				(3.54)		
$\sum R^{mo} - \Delta_4 pc$			−.007				−.031				−.049				−.037	
			(.796)				(3.15)				(4.23)				(3.98)	
$\sum R^{mo} - \Delta_4 pb$				−.005				−.028				−.044				−.037
				(.512)				(2.73)				(3.64)				(4.09)
coin-m	.054	.053	.053	.052	.086	.071	.084	.070	.085	.070	.082	.067	.085	.079	.090	.083
	(2.17)	(2.19)	(2.19)	(2.23)	(2.90)	(2.70)	(2.86)	(2.64)	(4.92)	(4.21)	(4.41)	(3.79)	(4.43)	(4.05)	(4.47)	(4.09)
$\overline{R}^2$	.877	.877	.877	.877	.885	.882	.884	.881	.892	.887	.890	.886	.893	.891	.894	.892
DW	2.04	2.05	2.04	2.04	2.06	2.08	2.05	2.06	2.16	2.17	2.14	2.14	2.06	2.08	2.05	2.07
LMAR(1-2)	.200	.221	.191	.206	.250	.354	.187	.280	1.04	1.07	.817	.833	.599	.741	.573	.696
	(.819)	(.802)	(.826)	(.814)	(.780)	(.703)	(.829)	(.758)	(.360)	(.349)	(.446)	(.439)	(.552)	(.480)	(.567)	(.502)
J.-B.	1.88	2.34	2.18	2.70	2.67	3.05	2.42	2.87	3.42	3.66	3.14	3.50	2.93	2.77	2.08	2.06
	(.390)	(.311)	(.336)	(.260)	(.263)	(.217)	(.299)	(.238)	(.181)	(.160)	(.208)	(.174)	(.231)	(.251)	(.354)	(.357)

ygap: Difference between log GDP and trend of log GDP, estimated by the expanded exponential smoothing filter. coin-m: Residuals of the estimated long-run money demand function using Divisia M3 or simple-sum M3. Estimation period: 1982Q1-2000Q4 for equations using Divisia M3 and 1983Q2-2000Q4 for equations using simple-sum M3. Further comments see Table 12.

the assessment of the test results. This is found for all analysed monetary aggregates, especially for DM^7. The problems posed by the assumption of normal distributed residuals decrease if the EES filter is applied. For this output gap the normality hypothesis is sometimes rejected at the 5 per cent significance level (see Table 12 continued). The coefficients of money changes are highly significant.

Table 13 exhibits the results of the IS-curve estimates, including excess liquidity. There are no severe problems with residual assumptions. The interest rate coefficients are negative and partly significant. The coefficient of the cointegration relation is positive, as expected. However, it is not significant for M3 and DM^1. Using DM^7, the estimated t-values are greater than 3.2, bearing in mind that the coefficient is not t-distributed but follows a non-standard distribution, like the Dickey-Fuller test statistic. Nevertheless, it seems sensible to conclude that the coefficient is significantly different from zero, at least at the 10 per cent level, since only one long-run coefficient is estimated.

The results presented are in line with evidence presented by Stracca (2001a) for DM^2. Overall, the outcome of this estimation supports the findings of Nelson (2001), that money enters significantly in the IS equation. It seems that the Divisia aggregate contains useful information for the policy-maker, which is not found in the real interest rate, on output. Moreover, the paper of Coenen, Levin and Wieland (2001) notes that real output data is often and substantially revised in the euro area over a period of up to nine periods. They show that especially money demand shocks calculated with simple-sum M3 contain information about the true level of output.

5.2 The P-Star approach

The long-run relationship between money and prices is based on the quantity equation

$$P \times Y = M \times V, \tag{44}$$

where P is the price level, Y is real output, M is the money supply, and V is the velocity of money. Owing to the definitions of the variables, the relationship in (44) is an identity.

By making two simplifying assumptions, the quantity equation becomes a theory of the cause of inflation. First, the velocity of money is regarded as depending on the institutional structure of the payments system. Since this system might be changed slowly over time, it is often suggested to treat V as being constant. If, second, output is exogenous for money and prices, changes in money must be reflected in changing prices. In a growing economy, Y may increase at some steady rate, thereby (partially) absorbing money growth. Furthermore, invariance of the velocity of money is a strong assumption, which should be tested empirically. As long as output and velocity are in equilibrium, however, equation (44) defines the equilibrium price level

$$P^* = (M/Y^*)V^*, \tag{45}$$

where equilibrium values are indicated with an asterisk (*). P^* aims to measure the price level to be obtained at actual money holdings if production and velocity are in equilibrium. If P and P^* are nonstationary and cointegrated, and the actual price level is below its equilibrium, a future acceleration of inflation can be expected (Hallman et al. 1991).

The equilibrium price level is not directly observable. To calculate P^*, empirical estimates of potential production and trend velocity are required. Potential output is often estimated by statistical methods (like those in chapter 5), but it is not apparent how to obtain trend velocity. If log velocity ($v_t = \ln V_t$) fluctuates randomly over time around a constant term, it becomes $v_t = v_0 + \epsilon_t$. If ϵ_t is a stationary zero mean process, the equilibrium level of log velocity is $v^* = v_0$. In some countries, however, velocities of monetary aggregates have exhibited a marked downward trend in the past. Orphanides and Porter (1998) assume a broken deterministic trend. Gottschalk and Bröck (2000) present different variants for the euro area data, whereas Scheide and Trabandt (2000) apply the Hodrick-Prescott filter. Rather than adopting a statistical method, Tödter and Reimers (1994) propose incorporating a stochastic trend of velocity if real money demand is income elastic ($\beta_1 > 1$)

$$m_t - p_t = \beta_0 + \beta_1 y_t + z_t, \tag{46}$$

where y_t is the log of real income (GDP), β_0 is a constant term and β_1 is the long-run income elasticity of money demand. If z_t is a stationary stochastic process with zero mean, equation (46) describes a cointegration relationship. King and Watson (1997) refer to (46) as being a monetary equilibrium condition. Contrary to this long-run relationship a short run dynamic money demand equation would have to take both lagged adjustment as well as interest rates into account.

Combining (44) and (46) yields the following expression for velocity:

$$v_t = -\beta_0 + (1 - \beta_1)y_t - z_t. \tag{47}$$

This suggests measuring trend velocity as

$$\begin{aligned} v_t^* &= -\beta_0 + (1 - \beta_1)y_t^* \\ &= v_0 + (1 - \beta_1)y_t^*. \end{aligned} \tag{48}$$

For $\beta_1 = 1$, this approach encompasses the stationary velocity case. If $\beta_1 > 1$ a declining trend in velocity is induced as long as potential output is growing.

Substituting (48) into the definition of P^* in (45), we end up with the following measure of equilibrium prices;

$$p_t^* = m_t - \beta_1 y_t^* + v_0. \tag{49}$$

The price gap is defined as

$$p_t^* - p_t = m_t - \beta_1 y_t^* + v_0 - (m_t - \beta_1 y_t + v_0) = \beta_1(y_t - y_t^*).$$

On the assumption of $\beta_1 = 1$. the price gap is the output gap. In such a case, the P*-approach is identical to the Phillips-curve approach. For $\beta_1 > 1$, the price gap additionally contains the velocity gap and accounts for the disequilibrium in the money market.

The investigation is conducted for two price measures. In line with the money demand analysis, in which real GDP approximates the transaction variable, the deflator of GDP (PB) is used. In contrast, the ECB defines price stability with respect to the increase in the harmonised index of consumer prices (PC). The development of both series and their annual growth rates are shown in Figure 10. The trend of the series seems to be identical. However, the inflation rate, measured as annual growth rate, is more smoothed for the PC than for the PB, whereas the standard deviations is greater for the growth rate of PC (.0346) than for the rate of PB (.0271). At the end of the sample, the changes in PC are higher than the changes in PB. The income elasticity β_1 is estimated by the Engle-Granger approach (see Engle and Granger, 1987).

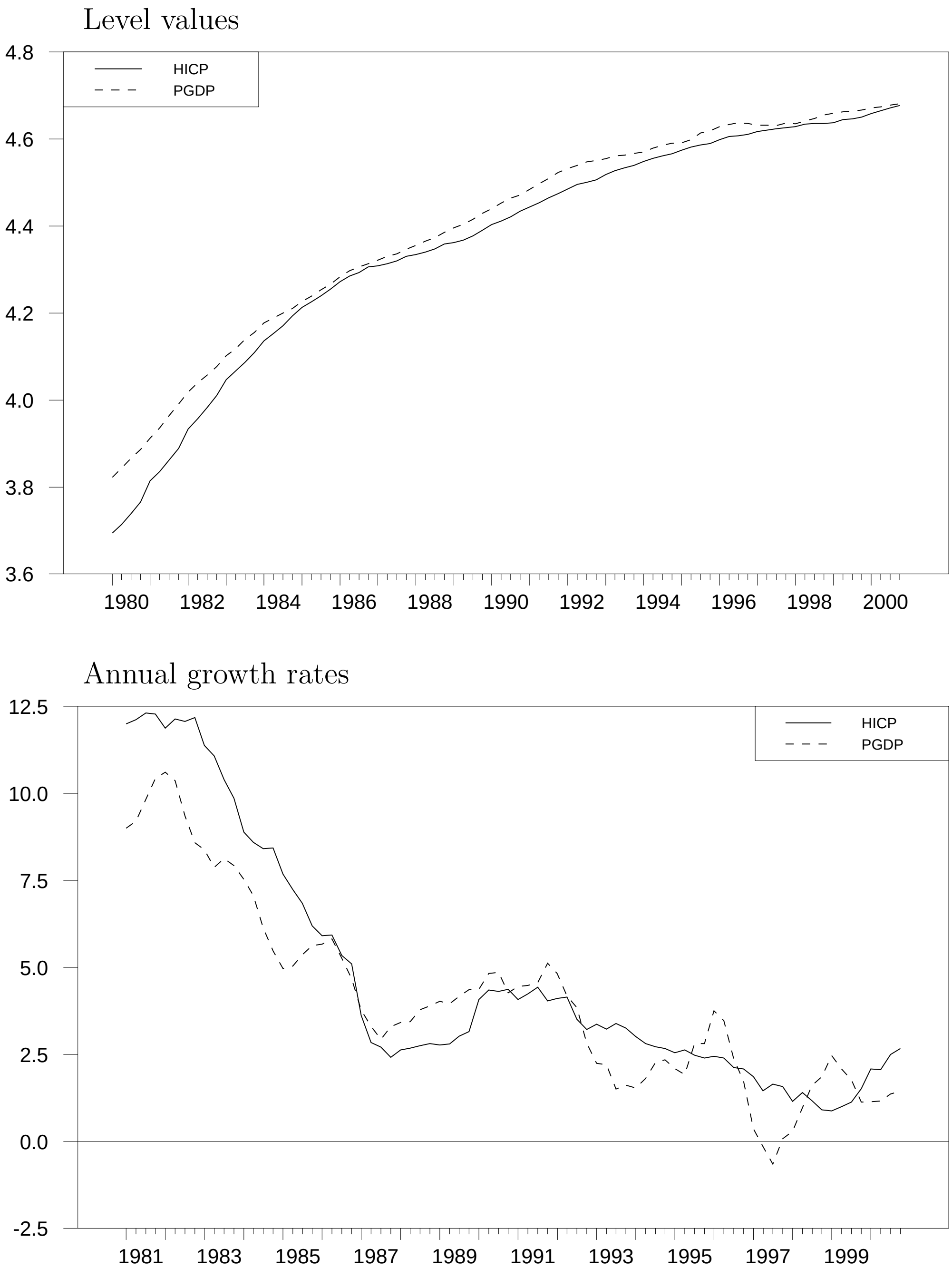

Figure 10: Levels of log price indices HICP (harmonised index of consumer prices) and PGDP (deflator of GDP), 1980-2000 (upper panel); Annual growth rates of the price indices in per cent, 1981-2000 (lower panel).

Table 14: Estimates of inflation equations with the output gap

	GDP deflator		HICP	
Method	HP	EES	HP	EES
Variable	Δp	Δp	Δp	Δp
$y_{t-1} - y^{\star}_{t-1}$	.027 (.611)	.038 (1.23)	.050 (1.96)	.052 (2.23)
Δp_{t-1}	.360 (3.85)	.349 (3.48)	.367 (4.78)	.332 (3.40)
Δp_{t-2}	.240 (2.49)	.301 (2.63)		
Δp_{t-3}			.102 (1.96)	.105 (1.99)
Δp_{t-4}	.240 (2.23)	.257 (2.70)	.686 (10.8)	.690 (14.4)
Δp_{t-5}			−.223 (3.48)	−.191 (2.29)
Δp^{oil}	−.007 (4.28)	−.007 (2.66)		
DUM871			−.012 (27.2)	−.011 (4.73)
$\overline{R}^2$	.717	.722	.917	.921
DW	1.97	1.98	1.93	1.98
LMAR(1-2)	.527 (.593)	.339 (.713)	.033 (.967)	.206 (.815)
J.-B.	.265 (.876)	.707 (.702)	2.21 (.331)	.775 (.679)

HP: using Hodrick-Prescott-filter; ESS: using expanded exponential smoothing filter. DUM871: Dummy variable is unity in 1987Q1 and zero elsewhere. Heteroskedasticity consistent covariance estimated t-values in parentheses. LMAR(1-2): Lagrange-multiplier test of autocorrelation of 1 and 2 lags. J.-B.: Jarque-Bera-test of normality. Diagnostic statistics have p-value in parentheses. Estimation period: 1982Q1-2000Q4.

The following estimates are obtained:

$$
\begin{aligned}
pb &= m3 + 7.020 - 1.477y \\
pb &= dm^1 + 5.916 - 1.312y \\
pb &= dm^3 + 6.572 - 1.409y \\
pb &= dm^7 + 5.151 - 1.203y \\
pc &= m3 + 6.109 - 1.353y \\
pc &= dm^1 + 5.005 - 1.188y \\
pc &= dm^3 + 5.661 - 1.285y \\
pc &= dm^7 + 4.240 - 1.080y.
\end{aligned}
$$

In all cases, the income elasticity is greater than unity. Potential output is calculated by the Hodrick-Prescott filter and the expanded exponential smoothing filter, as above. Moreover, the estimated inflation equation accounts for the oil price effects

$$\Delta p_t = \delta_0 + \delta_1(p_{t-1} - p^{\star}_{t-1}) + \delta_2 \Delta p^{\star}_t + \delta_3(L)\Delta p_{t-1} + \delta_4(L)p^{oil}_t + u_t,$$

where p^{oil}_t is the oil price, which is converted in euro using the current US-Dollar ecu exchange rates. The lag order used is $p = 6$. Coefficients of lagged inflation rates are set stepwise at zero if their estimated t-values are small in absolute terms. In order to have a reference result, a traditional Phillips curve is estimated. The results are presented in Table 14. The statistics of the diagnostic tests indicate no problems with the underlying residual assumptions. Regardless of the potential output estimate used, the output gap is not significant for the inflation rate measured by the GDP deflator. In contrast, it is significant at the 10 per cent level for the equation of PC inflation rate. These equations include an impulse dummy owing to a realignment in the EMS and drastic GDP changes in Italy. The dummy is unity in 1987Q1 and zero elsewhere.

Table 15 presents the results for the price gap approach. The first part includes the results for the GDP deflator the second part the estimates for consumer prices. The diagnostic tests analysed give no hints of problems with the residual assumptions. Most of the inflation rate measured according to the GDP deflator is explained by its own lags. The price gap is not significant for the GDP deflator regardless of the monetary aggregate used and the potential output estimate (see Table 15). The results are more mixed if the inflation rate of the HICP is examined (see Table 15 continued). The R^2 is high, over 0.85. If M3 or DM^3 are used to calculate the equilibrium price level, the price gap is not significant since a cointegrating relationship is specified. This result contradicts evidence given by Altimari (2001). He finds that M3 and equilibrium price level involving M3 helps to predict inflation rates in the euro area. The results change if DM^1 or DM^7 is considered. In these cases, the price gap is significant.

Hence, the conclusion may be drawn that some price gaps help to predict the inflation rate measured by the HICP. This evidence may be interpreted as a further indication that it is not wise to discard the information contained in the Divisia monetary aggregate.

Table 15: Estimates of inflation functions using P-star approach for GDP deflator

Method	HP	EES	HP	EES	HP	EES	HP	EES
Variable	M3		DM^1		DM^3		DM^7	
$p^*_{t-1} - p_{t-1}$	.010	.016	.020	.023	.009	.014	.045	.043
	(.681)	(1.10)	(1.55)	(1.96)	(.991)	(1.53)	(2.72)	(2.67)
Δp^*_t	.177	.182	.175	.169			.104	.103
	(2.41)	(2.75)	(3.02)	(3.24)			(2.33)	(2.54)
Δp_{t-1}	.313	.310	.315	.313	.365	.355	.317	.313
	(3.37)	(3.39)	(3.48)	(3.55)	(3.69)	(3.49)	(3.36)	(3.48)
Δp_{t-2}	.259	.260	.260	.263	.312	.306	.293	.291
	(2.15)	(2.19)	(2.26)	(2.28)	(2.63)	(2.53)	(2.49)	(2.44)
Δp_{t-4}	.182	.188	.196	.207	.233	.225	.227	.239
	(1.81)	(1.86)	(1.98)	(2.08)	(2.27)	(2.18)	(2.44)	(2.51)
Δp^{oil}_t	−.008	−.008	−.008	−.009	−.007	−.008	−.009	−.009
	(4.57)	(4.85)	(4.75)	(5.15)	(3.81)	(4.38)	(4.87)	(5.20)
$\overline{R}^2$	.730	.734	.739	.742	.719	.719	.743	.746
DW	1.98	1.99	1.97	1.98	1.97	1.99	2.01	2.03
LMAR(1-2)	.477	.390	.236	.196	.565	.431	.117	.106
	(.641)	(.679)	(.791)	(.823)	(.571)	(.651)	(.890)	(.900)
J.-B.	.548	.940	1.03	1.75	.224	.391	2.26	3.57
	(.760)	(.625)	(.596)	(.417)	(.894)	(.822)	(.323)	(.168)

P-star variable is constructed using the income elasticity estimates of an Engle-Granger regression and potential output variable. HP: using Hodrick-Prescott-filter; ESS: using expanded exponential smoothing filter. DUM871: Dummy variable is unity in 1987Q1 and zero elsewhere. Heteroskedasticity consistent covariance estimated t-values in parentheses. LMAR(1-2): Lagrange-multiplier test of autocorrelation of 1 and 2 lags. J.-B.: Jarque-Bera-test of normality. Diagnostic statistics have p-value in parentheses. Estimation period: 1982Q1-2000Q4.

6 Out-of-sample forecasts of prices and control errors of monetary aggregates

The money demand, IS-curve and price equations estimated in the previous sections empirically establish a link between monetary aggregates and prices. A monetary aggregate serving as an intermediate target for monetary policy must be controllable

Table 15 continued: Estimates of inflation functions using P-star approach for HICP

Method	HP	EES	HP	EES	HP	EES	HP	EES
Variable	M3		DM^1		DM^3		DM^7	
$p^*_{t-1} - p_{t-1}$	.012 (1.38)	.017 (2.16)	.019 (2.36)	.021 (3.23)	.003 (.414)	.006 (.869)	.032 (4.56)	.032 (5.15)
Δp^*_t	.153 (3.41)	.140 (3.53)	.129 (3.20)	.122 (3.19)	.077 (2.96)	.080 (3.24)	.110 (2.65)	.104 (2.73)
Δp_{t-1}	.111 (2.26)	.116 (2.54)	.127 (2.53)	.129 (2.74)	.346 (4.08)	.332 (3.76)	.136 (2.76)	.137 (2.99)
Δp_{t-4}	.671 (13.9)	.680 (14.3)	.690 (13.3)	.699 (14.0)	.714 (16.7)	.715 (16.6)	.700 (13.9)	.710 (15.0)
Δp_{t-5}					−.193 (2.97)	−.180 (2.67)		
Δp^{oil}_{t-3}	.004 (2.51)	.004 (2.34)	.005 (2.44)	.004 (2.33)	.004 (2.21)	.004 (2.11)	.004 (1.96)	.004 (1.91)
Dum871	−.009 (10.8)	−.009 (11.3)	−.009 (11.5)	−.009 (11.8)	−.011 (11.9)	−.011 (11.8)	−.010 (13.1)	−.010 (12.7)
$\overline{R}^2$	.923	.926	.924	.927	.922	.923	.929	.932
DW	1.95	1.99	1.90	1.96	2.13	2.13	1.93	2.01
LMAR(1-2)	.293 (.757)	.378 (.687)	.217 (.806)	.317 (.730)	1.01 (.369)	1.01 (.371)	.272 (.763)	.557 (.576)
J.-B.	.307 (.858)	.403 (.817)	1.06 (.588)	.954 (.621)	.551 (.759)	.689 (.708)	1.57 (.457)	1.21 (.546)

P-star variable is constructed using the income elasticity estimates of an Engle-Granger regression and potential output variable. HP: using Hodrick-Prescott-filter; ESS: using expanded exponential smoothing filter. DUM871: Dummy variable is unity in 1987Q1 and zero elsewhere. Heteroskedasticity consistent covariance estimated t-values in parentheses. LMAR(1-2): Lagrange-multiplier test of autocorrelation of 1 and 2 lags. J.-B.: Jarque-Bera-test of normality. Diagnostic statistics have p-value in parentheses. Estimation period: 1982Q1-2000Q4.

by monetary policy instruments. In this section, following Herrmann et al. (2000), we take the controllability of the monetary aggregates into account. The monetary framework developed so far may be summarised by the following equations

$$
\begin{aligned}
m - p &= f_1(y, R - r_o) + \eta_1 \quad \text{for simple-sum M3 or} \\
m - p &= \tilde{f}_1(y, pd) + \eta_1 \quad \text{for Divisia aggregates} \\
y - y^* &= f_2(y_{-1} - y^*_{-1}, m_{-1} - p_{-1}, r^{\text{real}}) + \eta_2 \\
p^* &= m - \hat{\beta}_1 y^* \qquad (50) \\
\Delta p &= f_3(p^* - p, \Delta p, \Delta p^*, \Delta p^{oil}) + u. \qquad (51)
\end{aligned}
$$

The ECB controls the target variable mainly with the aid of its interest rate policy. As a representative central bank interest rate, the ECB's rate for open market transactions of main refinancing operations (i_cen) is used. The demand for real money does not depend directly on this instrument, but rather on opportunity costs. In the case of M3 it is modelled as a function of the difference between a long-run benchmark rate and the own interest rate: $R - r_o$. Given the evidence in Brand and Cassola (2000), this spread is approximated by the long-run rate. For Divisia aggregates, opportunity costs are approximated by the corresponding price dual (pd), which is affected by the long-run rate (R) and interest rates of the components included that are approximated by the money market rate (R^{mo}). Hence, the interest rate link is estimated by simple dynamic term structure equations of the form

$$R = f_4(R_{-1}, i\_cen) + \eta_4$$

and, in addition, for Divisia aggregates by

$$\begin{aligned} R^{mo} &= f_5(R^{mo}_{-1}, i\_cen) + \eta_5 \\ pd &= f_6(R, R^{mo}) + \eta_6. \end{aligned}$$

The residual terms η_j for $j = 1, 2, 4, 5, 6$ contribute to the control error, while the residual term u in the inflation equation is called the projection error of the process. Even if the ECB could control its intermediate target perfectly, and even if it had accurate forecasts of the exogenous variable y^* of the process, it would not be able to control the rate of inflation perfectly because of the projection error. On the other hand, if, for example, DM^7 had the closest relationship to the rate of inflation but was controllable only with large errors, one of the other aggregates might perform better because of a smaller control error.

To investigate the whole process, we calculate a series of stepwise forecasts using the money demand and inflation equations for the different monetary aggregates, together with interest rate equations and the output gap function. To be in line with the estimated dynamic money demand functions of Chapter 4, the variables of the cointegrating vector are unrestrictedly included in the real money equations. It should be noted that the results of this exercise are conditional on the exogenous variables potential output and the oil price change. Moreover, using historical values of the interest rate instrument disregards the problem that the ECB would have set its rates differently if it had worked before 1999.

The out-of-sample forecasts are computed with a so-called recursive regression method (see McCracken, 1999). A recursive estimation of the system yields a series of out-of-sample forecasts for different forecasting horizons $k = 1, \cdots, 8$. The coefficients are computed over the period 1982Q1 to 1993Q4. Using these coefficients, the forecasts are determined. The forecast errors $\hat{e}_{t+k}$ are the difference between the forecast of the prices and the historical values. Then, the sample is extended by one period ahead and the equations are re-estimated to calculate the forecasts again. This procedure is continued until the end of the available data.

The projection error is determined by assuming that the monetary aggregates are exogenous. The forecasts are calculated by the equations (50) and (51), since the monetary aggregate affects the P^*-variable. This variable influences the inflation rate. This approach is denoted as the perfectly controlled money approach.

The benchmark approach is a restricted inflation equation

$$\Delta p = g_1(L)\Delta p + g_2(L)\Delta p^{oil} + \eta \,. \tag{52}$$

without the price gap and changes in P^*.

Since a complex system is used, it seems worthwhile to reduce the system in such a way that only the inflation equation is investigated, which additionally includes the lagged change in a money variable (see Baltensperger et al., 2001). The money variables are exogenous for this approach.

The accuracy of forecasts can be judged by various statistics about the forecast errors. In this study the root mean square forecast errors are presented. The mean absolute forecast errors point in the same direction. To assess the relative predictive accuracy of two forecasting models, different test statistics are suggested and analysed by Diebold and Mariano (1995). Their preferred test statistic is

$$\hat{d}_F = F^{-1/2} \frac{\sum_{t=T+1}^{S-k} (\hat{e}^2_{0,t+k} - \hat{e}^2_{1,t+k})}{\hat{\sigma}_F}, \tag{53}$$

where T denotes the length of estimation period, F is the length of the prediction period, hence $S = T + F$, $k \geq 1$ is the forecast horizon, $\hat{e}^2_{0,t+k}$ and $\hat{e}^2_{1,t+k}$ are squared forecast errors of the benchmark model and the alternative model using consistent estimators, and

$$\hat{\sigma}_F = \frac{1}{F}\sum_{t=T+1}^{S-k}(\hat{e}^2_{0,t+k}-\hat{e}^2_{1,t+k})^2 + \frac{2}{F}\sum_{j=1}^{l_F}\omega_j\sum_{t=T+1+j}^{S-k}(\hat{e}^2_{0,t+k}-\hat{e}^2_{1,t+k})(\hat{e}^2_{0,t+k-j}-\hat{e}^2_{1,t+k-j}),$$

where $\omega_j = 1-\frac{j}{l_F+1}$, $l_F = o(F^{1/4})$. The test statistic (53) is denoted the Diebold-Mariano (dm) test. The null of equal predictive ability is

$$H_0 = E(e^2_{0,t+k}-e^2_{1,t+k}) = 0,$$

while the alternative is

$$H_0 = E(e^2_{0,t+k}-e^2_{1,t+k}) \neq 0.$$

Under the null hypothesis, this statistic has an asymptotic standard normal distribution. Harvey, Leybourne and Newbold (1997, 1998) analyse the test statistic using an extensive Monte Carlo design, and find that the test has good size and fairly good power properties. Corradi, Swanson and Olivetti (2001) show that the asymptotic standard normal distribution property holds if cointegrated variables are investigated.

The longest interval for all forecasts is from 1994Q1 to 2000Q4, hence the maximum length of the forecast period is $F = 28$. The truncation parameter is $l_F = 2$. Table 16 (17) gives the results of the out-of-sample forecasts of the GDP deflator (HICP). The values RMSFE for model (52) increase for the GDP deflator if the forecasting horizon grows. The results of the other approaches are all given in relative values (RMSFE of the alternative approach divided by RMSFE of the benchmark model). Using the system approach with the different monetary aggregates it is apparent that the system with DM^1 outperforms the benchmark approach for some forecasting horizons. The other systems are worse than the benchmark equation for all horizons. The biggest forecast errors are obtained by the system with DM^7 for forecasting horizons $k = 2, \cdots, 8$. In some cases, the differences are significant at the 5% test level. Looking at the perfectly controlled money approach, it is apparent that the monetary aggregate DM^1 reduces the forecast errors of prices. This concept bears information for the

Table 16: Root mean squared forecast errors (RMSFE) for the GDP deflator and the relative results for systems using different monetary concepts and perfectly controlled money.

Hori-	System-approach					Perfectly controlled money			
zon	RMSFE	M3	DM^1	DM^3	DM^7	M3	DM^1	DM^3	DM^7
1	0.0040	1.02	0.99	1.00	1.01	1.02	0.99	1.00	1.01
2	0.0068	1.03	0.99	1.01	1.10	1.03	0.99	1.01	1.09
3	0.0102	1.04	1.00	1.03	1.19	1.04	0.99	1.02	1.17
4	0.0133	1.04+	0.99	1.03	1.26	1.04+	0.98	1.03	1.22
5	0.0169	1.06+	0.99	1.04	1.32+	1.05+	0.97	1.04	1.27
6	0.0205	1.08+	1.00	1.07	1.36+	1.07+	0.97	1.06	1.29
7	0.0233	1.09	1.01	1.09	1.46+	1.09+	0.97	1.08	1.32
8	0.0264	1.11	1.02	1.12	1.35+	1.10+	0.97	1.10	1.28

Ex ante root mean squared forecast errors for the period 1994Q1 - 2001Q1. Reference results of an autoregressive model for the inflation rate including the oil-price change $\Delta p_t = a_1 \Delta p_{t-1} + a_2 \Delta p_{t-2} + a_3 \Delta p_{t-4} + a_4 \Delta p_t^{oil} + u_t$. The sign '+' indicates that the difference between the benchmark model and the alternative model using the DM test is significant at the 5% level. Perfectly controlled money is realised by setting the different monetary aggregates exogenously.

future inflation rate. The differences in forecast errors between the system approach for monetary aggregate i and the perfectly controlled money approach for the same aggregate suggest that the control errors are small. Especially, for DM^7 are there large projection errors.

Turning to the HICP, the RMSFE of this index are markedly lower compared with the RMSFE of the GDP deflator (see Table 17, compared with Table 16). The inclusion of P^* variables does not reduce the forecasting errors in most cases. In contrast, using the DM^3 aggregate, this system outperforms the benchmark results for $k = 1, \cdots, 5$. Moreover, it is worth noting that the differences between squared forecast errors of the system approach and the benchmark model are not significant, owing to the relatively high standard errors of these differences. For example, the standard error used for the dm-test of $k = 3$ and the M3 system approach is six times higher compared with the perfectly controlled money approach. Looking at the results of this latter approach, it is apparent that the model including M3 reduces the forecast errors for all forecasting horizons. These results give hints of high control errors for this aggregate. Moreover, for DM^1 and DM^7 there exist striking control errors. In general, these results contradict

Table 17: Root mean squared forecast errors (RMSFE) for the HICP and the relative results for systems using different monetary concepts and perfectly controlled money.

Hori-	System-approach					Perfectly controlled money			
zon	RMSFE	M3	DM^1	DM^3	DM^7	M3	DM^1	DM^3	DM^7
1	0.0021	2.58	2.02	0.95	1.69	0.88	0.97	0.97	0.94
2	0.0033	2.08	1.67	0.98	1.50	0.79	1.28	1.00	1.03
3	0.0044	1.80	1.48	0.91	1.39	0.71+	1.25	1.00	1.06
4	0.0060	1.56	1.35	0.93	1.29	0.71+	1.15	0.99	1.02
5	0.0079	1.75	1.45	0.97	1.42	0.73+	1.11	1.03	1.08
6	0.0088	1.72	1.49	1.04	1.53	0.79	1.18	1.12	1.31
7	0.0097	1.61	1.47	1.11	1.57	0.84	1.21	1.22	1.49+
8	0.0104	1.56	1.55	1.17	1.72	0.92	1.28	1.31+	1.69+

Ex ante root mean squared forecast errors for the period 1994Q1 - 2001Q1. Reference results of an autoregressive model for the inflation rate including oil-price change $\Delta p_t = a_1 \Delta p_{t-1} + a_2 \Delta p_{t-4} + a_3 \Delta p^{oil}_{t-3} + a_4 DUM871 + u_t$. The sign '+' indicates that the difference between the benchmark model and the alternative model using the DM test is significant at the 5% level. Perfectly controlled money is realised by setting the different monetary aggregates exogenously.

the findings of the previous chapter. In that chapter evidence is presented that the P-star variable constructed by DM^7 is important for explaining the inflation rate. This difference may be put down to the fact that the in-sample results may not be adapted for the out-of-sample period.

At the end, the results are presented for the inflation equations including only money changes (see Table 18). The RMSEF are given for equation (52). For the GDP deflator, the accounting for money changes reduces the forecast errors. The reductions are significant for M3 over all forecasting horizons. Turning to HICP for $k = 3, 4, 5$, the monetary aggregates decrease the RMSFE. M3 outperforms the multiplicative aggregates. However, the results are not significantly different from the results of the benchmark model. On the assumption of exogeneity, money changes help to reduce the forecasting errors.

Table 18: Root mean squared forecast errors (RMSFE) for the GDP deflator and the HICP and the relative results for inflation equations involving money growth rates.

Hori-	GDP deflator					HICP				
zon	RMSFE	M3	DM^1	DM^3	DM^7	RMSFE	M3	DM^1	DM^3	DM^7
1	0.0040	0.96+	0.97	0.96	0.97	0.0021	0.97	1.00	1.00	1.00
2	0.0068	0.94+	0.95+	0.96	0.95	0.0033	0.93	0.96	1.00	0.96
3	0.0102	0.92+	0.94+	0.96	0.94	0.0044	0.91	0.94	0.93	0.94
4	0.0133	0.91+	0.92+	0.95	0.93+	0.0060	0.90	0.94	0.95	0.94
5	0.0169	0.91+	0.92+	0.94	0.93	0.0079	0.94	0.95	0.98	0.96
6	0.0205	0.90+	0.91+	0.94	0.92	0.0088	0.98	0.99	1.03	0.99
7	0.0233	0.89+	0.90+	0.93	0.91	0.0097	1.01	1.02	1.06	1.01
8	0.0264	0.87+	0.88+	0.91	0.89	0.0104	1.10	1.10	1.12	1.06

Ex ante root mean squared forecast errors for the period 1994Q1 - 2001Q1. Reference results of models given in Tables 16 and 17. The sign '+' indicates that the difference between the benchmark model and the alternative model using the dm-test is significant at the 5% level. The inflation equations additionally include the lagged money growth rate.

7 Conclusion

This study analyses historical Divisia aggregates for the euro area. Because monetary components of different countries have to be used, it is necessary to discuss alternative aggregation schemes. From a historical point of view, it seems appropriate to account for exchange rate changes until December 1998. Theoretically, the transaction weighting of national Divisia aggregates (DM^7) is least sensitive to exchange rate variations. This aggregate should present the historical money development in the euro area best of all

The main part of the study is an empirical examination of different Divisia aggregates, compared with simple-sum M3. In this investigation the first result is that the GARP-test indicates that it is possible to exclude money market funds and repo-funds from the summing up if a less broad aggregate than M3 is to be monitored.

Looking at the estimates of money demand functions for all Divisia aggregates, reasonable long-run equations may be determined. The income elasticity is mostly greater than unity. The coefficients of the opportunity cost measure are negative. The dynamic equations are stable and have reasonable statistical properties. Moreover, the central bank seems to affect them in the expected direction. In this sense, the

Divisia aggregates are controllable However, this influence is not always significant. Unexpected central bank interest rate innovations have more influence on DM^3 than on DM^7 and DM^1. There is no effect on M3.

The IS-curve estimates document the information content of money for real output movements. Especially, DM^7 includes valuable information on the future development of output. The importance of money for the inflation process is not as clear-cut as expected. For the in-sample exercise, the P-star framework is adopted. Inflation is measured by the annual growth rate of the HICP and the GDP deflator. The results of the inflation equations show that the price gap coefficients have the expected sign. Nevertheless, the coefficients of the price gaps are not always significantly different from zero. Only for DM^7 are the coefficients significant. This indicates a stable long-run link between money and prices. Hence DM^7 dominates the other aggregates.

The last test is the out-of-sample-forecast performance of simple inflation equations (perfectly controlled money approach) compared with more complicated system approaches. In this examination, none of the monetary aggregates improves the forecast errors of the growth rate of the GDP deflator, whereas the control errors are small. The control errors are higher regarding the growth rates of the HICP, especially using M3. On the other hand, a perfectly controlled M3 helps to forecast this inflation rate. Moreover, if money growth rates are directly put into the inflation equation they often reduce the forecast errors.

In sum, none of the aggregates dominates the others regarding all issues. Nevertheless, DM^7 seems to have stronger connections with output gap and price changes. This may be explained by the fact that DM^7 is the aggregate that includes smaller exchange rate effects than the others. Moreover, Divisia aggregates stress the transaction issue, and exclude the wealth component. Since the exchange rate changes are less important in the period immediately before the start of European Monetary Union and do not exist after January 1, 1999, this argument is not weakened by the fact that M3 helps to forecast HICP in the examined period. In general, the paper supports the view that money should have an important role in conducting monetary policy in the euro area, and that the ECB should investigate the movement of a Divisia aggregate.

Appendix: Estimation of Potential Output

The estimation of potential output Y^* is conducted by statistical methods. A linear function $y_t = f(t)$ is characterised by the fact that its first differences Δy_t are constant and its second differences $\Delta^2 y_t$ are zero. The Hodrick-Prescott (HP) filter adopt the second form (see Hodrick & Prescott, 1997). It is assumed that the series y may be divided into a trend component $\hat{y}$ and cyclical component y^c

$$y_t = \hat{y}_t + y_t^c.$$

The HP filter may be the solution of the following object function:

$$Z := \underset{(\hat{y}_t)}{\text{Min}} \left(\frac{\lambda}{2} \sum_{t=2}^{T} ((\hat{y}_{t+1} - \hat{y}_t) - (\hat{y}_t - \hat{y}_{t-1}))^2 + \frac{1-\lambda}{2} \sum_{t=1}^{T} (\hat{y}_t - y_t)^2 \right)$$

It results in the following:

$$Y^* = H^{-1} Y$$

where

$$H = \frac{1}{1-\lambda} \begin{pmatrix} 1 & -2\lambda & \lambda & 0 & \cdots & 0 & 0 \\ -2\lambda & 1+4\lambda & -4\lambda & \lambda & \cdots & 0 & 0 \\ \lambda & -4\lambda & 1+5\lambda & -4\lambda & \cdots & 0 & 0 \\ \vdots & & & \ddots & \ddots & & \vdots \\ 0 & & & & & -4\lambda & \lambda \\ 0 & & & & & 1+4\lambda & -2\lambda \\ 0 & & & & & -2\lambda & 1 \end{pmatrix}$$

Except for the first and last two observations, the filter relation is:

$$\hat{y}_t = \frac{6\lambda}{1+r\lambda} \tilde{y}_t + \frac{1-\lambda}{1+r\lambda} y_t \quad t = 3, \cdots, T-2,$$

where

$$\tilde{y}_t = \frac{y_{t-2} + 4\hat{y}_{t-1} + 4\hat{y}_{t+1} - \hat{y}_{t+2}}{6}$$

Mohr (2001) discusses the structural breaks and the end-point problems posed by the HP filter as well as the choice of smoothing parameter λ. In the empirical literature the value of $\lambda = \frac{1600}{1+1600}$ is often used for quarterly data. Tödter (2001) presents calculations that this value implies a reference cycle of 8 to 9 years for a business cycle. He shows that a reference business cycle of 8 years implies a value of $\lambda = \frac{1410}{1+1410}$, which is close to the standard value. Pedersen (2001) argues that the HP filter with the standard value of $\lambda = \frac{1600}{1+1600}$ is in many cases less distorting than other filters.

Adopting the HP filter for monthly data, changes the adjustment parameter. According to Ravn and Uhlig (2001), the value should be $\lambda = \frac{129600}{1+129600}$ if the starting point is the standard value of $\lambda = \frac{1600}{1+1600}$.

In contrast to the Hodrick-Prescott filter, Tödter shows that extended exponential smoothing (EES) uses the assumption that the first difference of a series is constant. Following Tödter (2000a), the EES procedure is derived from the function:

$$Z := \underset{(\hat{y}_t, c_1)}{\text{Min}} \left(\frac{\lambda}{2} \sum_{t=2}^{T} (\hat{y}_t - \hat{y}_{t-1} - c_1)^2 + \frac{1-\lambda}{2} \sum_{t=1}^{T} (\hat{y}_t - y_t)^2 \right)$$

The first term reflects the smoothness of the filtered series and the second term gives the adjustment of the estimated series to the observed series. The first order conditions are determined by differencing the function to all $\hat{y}_t$ and c_1. The conditions imply that the intercept term c_1 may be determined by the following nonparametric estimate:

$$\hat{c}_1 = \frac{1}{T-1} \sum_{t=2}^{T} (\hat{y}_t - \hat{y}_{t-1}) = \frac{\hat{y}_T - \hat{y}_1}{T-1}$$

The filtered series is:

$$Y^* = A^{-1} Y$$

where

$$A = \frac{1}{1-\lambda} \begin{pmatrix} 1-\frac{\lambda}{T-1} & -\lambda & 0 & 0 & & & \cdots & \frac{\lambda}{T-1} \\ -\lambda & 1+\lambda & -\lambda & 0 & & & \cdots & 0 \\ 0 & -\lambda & 1+\lambda & -\lambda & & & \cdots & 0 \\ \vdots & & & \ddots & \ddots & & & \vdots \\ 0 & & & & 1+\lambda & -\lambda & & 0 \\ 0 & & & & -\lambda & 1+\lambda & & -\lambda \\ \frac{\lambda}{T-1} & 0 & & \cdots & 0 & -\lambda & & 1-\frac{\lambda}{T-1} \end{pmatrix}$$

The filter, which is denoted as extended exponential smoothing, is:

$$Y_t^* = \frac{2\lambda}{1+\lambda} \left(\frac{Y_{t-1}^* + Y_{t+1}^*}{2} \right) + \frac{1-\lambda}{1+\lambda} Y_t$$

for $t = 1, \cdots, T-1$ and λ smoothing parameter. Assuming that the EES is an approximation of an optimal filter for a reference cycle of 8 years, $\lambda = 132/133$ (see Tödter, 2000b).

References

Altimari, N. (2001), *Does Money Lead Inflation in the Euro Area?*, European Central Bank Working Paper Series, No. 63, Frankfurt am Main.

Baltensperger, E., T.J. Jordan, M.R. Savioz (2001), 'The Demand for M3 and Inflation Forecasts: An Empirical Analysis for Switzerland', *Weltwirtschaftliches Archiv*, vol. 137, pp. 244-272.

Barnett, W.A. (1978), 'The User Cost of Money', *Economics Letters*, vol. 1 , pp. 145-149.

Barnett, W.A. (1980), 'Economic Monetary Aggregates: An Application of Index Number and Aggregation Theory', *Journal of Econometrics*, vol. 14, pp. 11-48.

Barnett,W.A., D. Fisher, A. Serletis (1992), 'Consumer Theory and the Demand for Money', *Journal of Economic Literature*, vol. 30, pp. 2086-2119.

Barnett,W.A., Y. Liu, M. Jensen (2000), 'CAPM Risk Adjustment' W.A. Barnett, A. Serletis (eds.), The Theory of Monetary Aggregation, pp. 245-273.

Barnett,W.A., E.K. Offenbacher, P.A. Spindt (1984), 'The New Divisia Monetary Aggregates', *Journal of Political Economy*, vol. 92, pp. 687-710.

Barnett, W.A., M. Hinich, P. Yue (2000), 'Monetary Aggregation Theory Under Risk Aversion', in: W.A. Barnett, A. Serletis (eds.), The Theory of Monetary Aggregation, pp. 217-244.

Bauwens, L., J. Hunter (2001), 'Identification and Exogeneity in the Long-Run', Discussion Paper, Brunel University.

Bayoumi, T., P.B. Kenen (1993), 'How Useful is an EC-wide Monetary Aggregate as an Intermediate Target for Europe?', *Review of International Economics*, vol. 1, 209-220.

Beyer, A., J.A. Doornik, D.F. Hendry (2001), 'Constructing Historical Euro-Zone Data', *Economic Journal*, vol. 111, pp. 308-327.

Brand, C., N. Cassola (2000): *A Money Demand System for Euro Area M3*, Working Paper No. 39, European Central Bank, Frankfurt am Main.

Calza, A., D. Gerdesmeier, J. Levy (2001): *Euro Area Money Demand: Measuring the Opportunity Cost Appropriately*, IMF Working Paper No. 01/179, Washington.

Coenen, G., A. Levin, V. Wieland (2001): *Data Uncertainty and the Role of Money as an Information Variable for Monetary Policy*, Working Paper No. 84, European Central Bank, Frankfurt am Main.

Coenen, G., J.-L. Vega (1999): *The Demand for M3 in the Euro Area*, Working Paper No. 6, European Central Bank, Frankfurt am Main.

Corradi, V., N.R. Swanson, C. Olivetti (2001), *Predictive Ability With Cointegrated Variables*, Discussion Paper, Texas A&M University.

Dedola, L., E. Gaitto, L. Silipo (2001), *Money Demand in the Euro Area: Do National Differences Matter?*, Banca d'Italia, Temi di Discussione, No. 405.

Deutsche Bundesbank
(2001), 'Monetäre Entwicklung im Euro-Währungsgebiet seit Beginn der EWU', *Monthly Report*, vol. 53, June, pp. 41-58.

Diebold, F.X., R.S. Mariano (1995), 'Comparing Predictive Accuracy', *Journal of Business & Economic Statistics*, vol. 13, pp. 252-263.

Drake, L., A. Mullineux, J. Agung (1997), 'One Divisia Money for Europe?' *Applied Economics*, vol. 29, pp. 775-786.

ECB (1999), 'Euro Area Monetary Aggregates and Their Role in the Eurosystem's Monetary Policy Strategy', *Monthly Bulletin*, European Central Bank, vol. 1 (February), pp. 29-47.

ECB (2001), *The Monetary Policy of the ECB*, European Central Bank, Frankfurt am Main.

Engle, R.F., C.W.J. Granger (1987): Co-integration and Error Correction: Representation, Estimation and Testing, *Econometrica*, vol. 55, pp. 251-276.

Fase, M.M.G. (2000), Divisia Aggregates and the Demand for Money in Core EMU, in: M. Belongia & J.M. Binner (eds.) 'Divisia Monetary Aggregates, Right in Theory - Useful in Practice?', Palgrave, Houndmills, pp . 138-169.

Friedman, M. (1956), *The Quantity Theory of Money - a Restatement*, in Friedman, M. (ed.) Studies in the Quantity Theory of Money, University of Chicago Press, pp. 3-21.

Gaab, W., A. Mullineux (1996), *Monetary Aggregates and Monetary Policy in the UK and Germany*, D. Duwendag (ed.) *Finanzmärkte, Finanzinnovationen und Geldpolitik*, Duncker & Humblot: Berlin.

Gottschalk, J., S. Bröck (2000), 'Inflationsprognosen für den Euro–Raum: Wie gut sind P*–Modelle?', *Vierteljahrshefte zur Wirtschaftsforschung*, vol.69, pp. 69-89.

Hallman, J.J., R.D. Porter, D.H. Small (1991), 'Is the Price Level Tied to the M2 Monetary Aggregates in the Long Run?', *The American Economic Review*, 81, 841–858.

Hansen, H., S. Johansen (1999), 'Some Tests for Parameter Constancy in Cointegrated VAR-Models', *Econometric Journal*, vol. 2, pp. 306-333.

Hansen, H., K. Juselius (1995), 'CATS in RATS: Cointegration Analysis of Time Series', Discussion Paper, Estima, Evanston, IL.

Harvey, D.I., S.J. Leybourne, P. Newbold (1998), 'Tests for Forecast Encompassing', *Journal of Business & Economic Statistics*, vol. 16, pp. 254-259.

Harvey, D.I., S.J. Leybourne, P. Newbold (1997), 'Testing the Equality of Prediction Mean Squared Errors', *International Journal of Forecasting*, vol. 13, pp. 281-291.

Herrmann, H., H.-E. Reimers & K.-H. Toedter (2000), 'Weighted Monetary Aggregates for Germany', in: M. Belongia & J.M. Binner (eds.) 'Divisia Monetary Aggregates, Right in Theory - Useful in Practice?', Palgrave, Houndmills, pp. 79-100.

Hodrick, R., E. Prescott (1997), 'Post-War U.S. Business Cycles: An Empirical Investigation', *Journal of Money, Credit, and Banking*, vol. 29, pp. 1-16.

Johansen, S. (1988), 'Statistical Analysis of Cointegration Vectors', *Journal of Economic Dynamics and Control*, vol. 12, pp. 231-254.

Johansen, S. (1991), 'Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", *Econometrica*, vol. 59, pp. 1551-1580.

Johansen, S., K. Juselius (1992), 'Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and UIP for UK', *Journal of Econometrics*, vol. 53, pp. 211-244.

Johansen, S., K. Juselius (2001), 'Controlling Inflation in a Cointegrated Vector Autoregressive Model with an Application to US Data', EUI Working Paper ECO No. 2001/2, European University Institute, Badia Fiesolana, San Domenico.

King, R.G., M.W. Watson (1997), Testing Long-Run Neutrality, *Economic Quarterly, Federal Reserve Bank of Richmond*, vol. 83 (Summer), pp. 69–101.

Lütkepohl, H. (1991), *Introduction to Multiple Time Series Analysis*, Springer–Verlag: Berlin.

McCracken, M.W. (1999), *Asymptotics for Out of Sample Tests of Causality*, Discussion Paper, Louisiana State University.

Mohr, M. (2001), *Ein disaggregierter Ansatz zur Berechnung konjunkturbereinigter Budgetsalden für Deutschland: Methoden und Ergebnisse*, Diskussionspapier 13/01, Volkswirtschaftliches Forschungszentrum der Deutschen Bundesbank, Frankfurt am Main.

Nelson, E. (2001), *Direct Effects of Base Money on Aggregate Demand: Theory and Evidence*, CEPR Discussion Paper No. 2666.

Orphanides, A., R. Porter (1998), *P* Revisited: Money-Based Inflation Forecasts with a Changing Equilibrium Velocity*, Discussion Paper, Board of Governors of the Federal Reserve System, Washington.

Osterwald-Lenum, M. (1992), 'A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics', *Oxford Bulletin of Economics and Statistics*, vol. 54, pp. 461-472.

Pedersen, T.M. (2001): The Hodrick-Prescott filter, the Slutzky effect and the distortionary effect of filters, *Journal of Economic Dynamics & Control*, vol. 25, 1081-1101.

Ravn, M.O., H. Uhlig (2001), *On Adjusting the HP-filter for the Frequency of Observations*, Center for Economic Studies & Ifo Institute for Economic Research (CESifo) Working Paper No. 479.

Reimers, H.-E., K.-H. Tödter (1994), *Europäische Geldmengenaggregation*, mimeo, Deutsche Bundesbank, Frankfurt am Main.

Reischle, J. (2000), Der Divisia-Geldmengenindex: Eine Analyse seiner theoretischen Grundlagen, Europäische Hochschulschriften, Peter Lang Verlag, Frankfurt am Main.

Rotemberg, J.J., M. Woodford (1997), 'An Optimisation-based Econometric Framework for the Evaluation of Monetary Policy', in: B.S. Bernanke & J.J. Rotemberg (eds.) *NBER Macroeconomics Annual 1997*, MIT Press, pp. 297-346.

Rotemberg, J.J., M. Woodford (1999), 'Interest Rate Rules in an Estimated Sticky Price Model', in J.B. Taylor (ed.) *Monetary Policy Rules*, University of Chicago Press, pp. 57-119.

Rudebusch, G.D., L.E.O. Svensson (1999), *Policy Rules of Inflation Targeting, in J.B. Taylor (ed.)*, Monetary Policy Rules, University of Chicago Press, pp. 203-246.

Rudebusch, G.D., L.E.O. Svensson (2000), *Eurosystem monetary targeting: Lessons from US Data*, mimeo IIES.

Scharnagl, M. (1996), 'Monetary Aggregates with Special Reference to Structural Changes in the Financial Markets', Discussion Paper, Economic Research Group of the Deutsche Bundesbank, Frankfurt.

Scheide, J., M. Trabandt (2000), *Predicting Inflation in Euroland – The Pstar Approach*, Kiel Working Paper No. 1019, Kiel.

Spencer, P. (1995), 'Should the ECB Adopt a Divisia Monetary Aggregate?', Paper presented at the CEPR Frankfurt conference"What monetary policy for the ECB?", mimeo.

Stracca, L. (2001a), *Does Liquidity Matter? Properties of a Synthetic Divisia Monetary Aggregate in the Euro Area*, Working Paper No. 79, ECB, Frankfurt am Main.

Stracca, L. (2001b), *The Functional Form of the Demand for Euro Area M1*, Working Paper No. 51, ECB, Frankfurt am Main.

Tödter, K.-H. (2000a): *Erweiterte exponentielle Glättung als Instrument der Zeitreihenanalyse, I. Eine Alternative zum Hodrick-Prescott-Filter*, Deutsche Bundesbank, mimeo.

Tödter, K.-H. (2000b), *Erweiterte exponentielle Glättung als Instrument der Zeitreihenanalyse, II. Zur Wahl des Glättungsparameters*, Deutsche Bundesbank, mimeo.

Tödter, K.-H. (2001), *Erweiterte exponentielle Glättung als Instrument der Zeitreihenanalyse, III. Idealer Filter, Gewichtung und Revisionsverhalten*, Deutsche Bundesbank, mimeo.

Varian, H.R. (1982), 'The Nonparametric Approach to Demand Analysis', *Econometrica*, vol. 50, pp. 945-973.

Varian, H.R. (1983), 'The Non-parametric Tests of Consumer Behoviour', *Review of Economic Studies*, vol. 50, pp. 99-110.

Wesche, K. (1997), 'The Demand for Divisia Money in a Core Monetary Union', *Review of the Federal Reserve Bank of St. Louis*, vol. 79, September/October, pp.51-60.

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