Lemke, Wolfgang; Greiber, Claus

Working Paper
Money demand and macroeconomic uncertainty

Discussion paper Series 1 / Volkswirtschaftliches Forschungszentrum der Deutschen Bundesbank, No. 2005,26

Provided in Cooperation with:
Deutsche Bundesbank

Suggested Citation: Lemke, Wolfgang; Greiber, Claus (2005) : Money demand and macroeconomic uncertainty, Discussion paper Series 1 / Volkswirtschaftliches Forschungszentrum der Deutschen Bundesbank, No. 2005,26, Deutsche Bundesbank, Frankfurt a. M.

This Version is available at:
http://hdl.handle.net/10419/19611

Terms of use:
Documents in EconStor may be saved and copied for your personal and scholarly purposes.
You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.
If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

Standard-Nutzungsbedingungen:
Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.
Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.
Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.
Money demand and macroeconomic uncertainty

Claus Greiber
Wolfgang Lemke

Discussion Paper
Series 1: Economic Studies
No 26/2005

Discussion Papers represent the authors’ personal opinions and do not necessarily reflect the views of the Deutsche Bundesbank or its staff.
Abstract:

In this study we construct a measure of macroeconomic uncertainty from several observable economic indicators for the euro area. Indicator variables are based on financial market data, such as medium-term returns, loss and volatility measures but also come from surveys that capture business and consumer sentiment. From these we estimate the path of underlying macroeconomic uncertainty using an unobserved components model. Employing cointegration analysis it is demonstrated that the extracted measures of uncertainty help to explain the increase in euro area M3 over the period 2001 to 2004. Similar evidence can be found for US monetary aggregates.

JEL Classification: E41

Keywords: Money demand, Macroeconomic Uncertainty, Excess Liquidity
Non technical summary

Over the last three years M3 growth in the euro area has been constantly above the reference value. As a consequence, hitherto stable standard money demand models failed to explain the observed monetary developments and showed signs of instability in recent periods. One explanation for this phenomenon claims that an environment of increased macroeconomic uncertainty in conjunction with low asset yields has enhanced the preference for liquidity. However, with regard to economic analysis uncertainty is difficult to capture on a conceptual level as well as in terms of quantification and measurement. Most of the literature employs a particular observable indicator as a proxy for uncertainty. The restriction on such a particular variable may sometimes be deemed as somewhat arbitrary, yet. In contrast, our analysis treats macroeconomic uncertainty as an unobserved process which itself is reflected in several observable economic indicators. From these measurements we then estimate the path of the underlying macroeconomic uncertainty using an unobserved components model. Our indicator variables are mainly based on financial market data, such as medium-term returns, loss and volatility measures but also come from surveys that capture business and consumer sentiment. It is shown that the extracted measures of uncertainty help to explain the increase in euro area M3 over the period 2001 to 2004. In particular, a cointegrated money demand relationship can be established for samples that include these periods. The robustness of our approach is tested by applying it to US data. Accordingly, similar forces seemed to be at work in the US. Augmenting money demand specifications for certain US monetary aggregates by uncertainty measures enhances their empirical performance, too. Thus, the study supports the assessment that monetary growth in recent years was significantly influenced by portfolio shifts from risky to liquid assets which were triggered by an increased level of economic uncertainty.
Nicht technische Zusammenfassung

Contents

1 Introduction ................................................. 1

2 Conceptual Issues .......................................... 2
   2.1 Money in a CCAPM ...................................... 3
   2.2 Uncertainty ............................................. 5

3 Empirical Methods ......................................... 6
   3.1 Empirical Model and Estimation Approach .......... 6

4 Euro area .................................................. 10
   4.1 Variables and Data ...................................... 10
   4.2 Empirical Results ....................................... 14

5 The Case of the US ......................................... 19
   5.1 Monetary Measures for the US ......................... 19
   5.2 The US Uncertainty Model ............................. 20
   5.3 US Money Demand Functions ......................... 25

6 Conclusions ................................................. 31
List of Tables

1  M3 money demand, euro area, 1980Q1 - 2004Q4 .......................... 16
2  M2 money demand ................................................................. 26
3  M2M money demand, 2 factors - factor 1 only ............................ 27
4  M2M money demand, 1 factor ................................................... 28
5  MZM money demand, 2 factors, factor 1 only ............................. 29
List of Figures

1 Log of M3, log of real income, opportunity cost in levels, and opportunity cost in logs. .................................................. 11
2 Correlation between stock and bond returns, stock market loss, stock market volatility and stock market return. .................................................. 12
3 Consumer confidence and industry confidence. .................................................. 13
4 Estimated factors representing uncertainty .................................................. 15
5 Excess liquidity (EC_{t-1}) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in levels. .................. 18
6 Excess liquidity (EC_{t-1}) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in logs. .................. 18
7 Measurements, US unobserved components model .................................. 22
8 Uncertainty factors, two-factor model .................................................. 24
9 Uncertainty factor, one-factor model .................................................. 24
10 Monetary overhang, M2M .................................................. 30
11 Monetary overhang, MZM .................................................. 30
Money Demand and Macroeconomic Uncertainty*

1 Introduction

In recent years, money growth in the euro area has been stronger than what could be expected against the background of real GDP and interest rate movements. According to official statements of the ECB this should be attributed to portfolio-shifts from risky to safe and liquid assets which were initially caused by the decline in stock markets and feded afterwards by a general sentiment of geopolitical and economic uncertainty:

"Nonetheless, M3 growth remains resilient. It appears that the reversal of past portfolio shifts is proceeding more slowly than would have been expected on the basis of historical regularities. This may reflect an increased risk aversion of households and firms, given the stock market losses they experienced between 2000 and the spring of 2003. In addition, the low level of interest rates continues to support monetary expansion, especially of the most liquid assets included in the narrow aggregate M1."

This assessment is reflected by the fact that standard money demand models have not been able to explain recent monetary developments anymore. As a consequence, there is evidence for a current instability of money demand functions in the euro area. Using the end-of-sample break point test by Andrews and Kim (2003), Carstensen (2004) establishes a break in usual “workhorse” money demand specifications like Calza, Gerdesmeier, and Levy (2001). On the contrary, Bruggeman, Donati, and Warne (2003) show that Nyblom-type stability tests might be flawed by small-sample biases which could lead to a spurious detection of break points in euro area money demand functions. Though this conflicting evidence might also be due to slightly different definitions of stability, a final assessment of this issue does not seem to be possible, yet. However, the current degree of “unexplainedness” of monetary developments in the euro area seems to be significant enough to make it worthwhile to explore the presumed portfolio shift story further.

In connection with that, some empirical approaches show that augmenting standard money demand models by additional variables help to improve the explanation of monetary developments. Carpenter and Lange (2002), Kontolemis (2002) and Carstensen

*The views expressed here are those of the authors and not necessarily those of the Deutsche Bundesbank. We thank seminar participants at Deutsche Bundesbank and the Österreichische Nationalbank.

*Introductory statement to the press conference Jean-Claude Trichet, President of the ECB, Lucas Papademos, Vice President of the ECB, 2 September 2004.
(2004) establish an empirical link between money demand and stock price as well as related measures. However, those approaches are rather eclectic and sometimes lack a well defined reasoning of the nature of portfolio shifts. A more theory-based framework for augmenting the standard money demand specification is presented by Choi and Oh (2003) and Attah-Mensah (2004). Those studies specify an exogenous preference shift that influences the utility derived from holdings of liquid assets. This finally motivates empirical money demand models in which an additional variable (alongside output and interest rates) represents this preference term.

Our approach follows this literature and attempts to reduce the degree of “eclecticism”. We assume (1) that the long-run variation in the preference for liquidity is determined by the level of “macroeconomic uncertainty” and (2) that this uncertainty can be captured to a large extent by one or two unobserved factors. In doing so, we understand uncertainty as a fairly broad concept capturing the bundle of forces that have led to a shift in the preference for liquidity. These include high suffered stock market losses, high experienced volatility and unsettling geopolitical events. Accordingly, the underlying factors are extracted from ex post indicators representing mainly financial market developments such as assets returns, loss and volatility measures as well as corporate bond spreads. Additionally, confidence indicators based on surveys are employed.

The empirical results show that our approach is successful in explaining the recent abundance of liquidity in the euro area. It is demonstrated that augmenting money demand functions by indicators of macroeconomic uncertainty establishes cointegration among money holdings and its determinants for the euro area. Moreover, uncertainty measures similar to those employed for the euro area tend to improve the explanation of some monetary aggregates in the US.

Our study is organised as follows. In section 2 we present a theoretical background for our empirical model and discuss the notion of uncertainty. Section 3 develops the empirical model and 4 estimates it for the euro area. As a check of robustness, section 5 applies the approach to the US. The final section concludes.

2 Conceptual Issues

This section presents a conceptual framework in which our analysis is carried out. The first part builds on a consumption-based capital asset pricing model (CCAPM) which serves as a general framework for the specification of an empirical money demand function. Second, we discuss the meaning of “uncertainty” in the context of our study.
2.1 Money in a CCAPM

In this theoretical illustration, we use a standard money-in-the-utility-approach (MIU) with an additive separable utility function. In this setting a representative agent can choose between money, safe bonds and a risky asset to transfer purchasing power from today to future periods.\(^2\)

Consider a consumer who maximises his lifetime utility depending on real consumption and money balances

\[
\max E_0 \left[ \sum_{t=0}^{\infty} \beta^t u \left( C_t, \frac{M_t}{P_t} \right) \right]
\]

over his choices for real consumption \(C_t\) and nominal money \(M_t\).

Each period the representative agent receives real income \(Y_t\), real dividends \(d_t\) and gross real returns \(R_t\) from equity \(Q_{t-1}\) and one-period bond holdings \(B_{t-1}\) respectively. Moreover, he sells real equities at the real price \(q_t\) while bond and money holdings trade at the price of one. This income is spent on consumption, investment in equities, bonds and money holdings. Hence, the representative consumer faces the budget constraint

\[
C_t + q_t Q_t + \frac{B_t}{P_t} + \frac{M_t}{P_t} = (q_t + d_t) Q_{t-1} + R_{t-1} \frac{B_{t-1}}{P_t} + \frac{M_{t-1}}{P_t} + Y_t.
\]

Defining the stochastic discount factor as

\[
\mathcal{M}_{t+1} := \beta \frac{u_C(C_{t+1}, M_{t+1}/P_{t+1})}{u_C(C_t, M_t/P_t)},
\]

the first-order optimality conditions of the resulting maximisation problem can be written as

\[
E_t \left[ \mathcal{M}_{t+1} R_t \right] = 1,
\]

\[
E_t \left[ \mathcal{M}_{t+1} q_{t+1} + d_{t+1} \right] = 1
\]

and

\[
E_t \left[ \mathcal{M}_{t+1} \frac{P_t}{P_{t+1}} \right] + \frac{u_M(P_t, M_t/P_t)}{u_C(C_t, M_t/P_t)} = 1.
\]

Assuming a Fisher relationship

\[
R_t = \frac{P_t}{P_{t+1}} (1 + i_t),
\]

where the nominal interest rate from \(t\) to \(t+1\) is given by \(i_t\), (2.4) can be written as

\[
E_t \left[ \mathcal{M}_{t+1} \frac{P_t}{P_{t+1}} \right] = \frac{1}{1 + i_t}.
\]

Inserting into (2.5) and rearranging one obtains

\[
\frac{u_{M/P}(C_t, M_t/P_t)}{u_C(C_t, M_t/P_t)} = \frac{i_t}{1 + i_t}.
\]  

(2.9)

In order to obtain explicit solutions, we assume a CRRA utility function of the type

\[
u(C_t, M_t/P_t) = \frac{C_t^{1-\sigma} - 1}{1 - \sigma} + b_t^\sigma (M_t/P_t)^{1-\gamma} - 1, \]

(2.10)

where \(\sigma\) and \(\gamma\) represent coefficients of relative risk aversion with respect to consumption and real balances. The random process \(b_t\) is a preference shift variable with \(\delta\) governing its impact on utility derived from money holdings.

Combining (2.6), (2.9), (2.10) and solving for real money holdings leads to a money demand function of the form

\[
\ln(M_t/P_t) = \frac{\sigma}{\gamma} \ln(C_t) - \frac{1}{\gamma} \ln\left(\frac{i_t}{1 + i_t}\right) + \frac{\delta}{\gamma} \ln(b_t).
\]

(2.11)

This equation states that real balances are determined by consumption (being equal to output in equilibrium) and the safe interest rate. Additionally, the liquidity preference shift variable \(b_t\) can have an impact on money holdings. Generally, it is allowed to be highly persistent. This approach serves as the general framework for our empirical investigations.

Note that, with a view to empirical work, this framework does not provide a direct and additional influence of risky asset returns \(q_t\) on money demand if the safe interest rate is included. This is because the system of first order conditions implies that either (2.4) or (2.5) is redundant for the money demand equation. It can be written either in terms of the safe interest rate \(i_t\) or the expected yield on risky assets \(\frac{q_{t+1} + d_{t+1}}{q_t}\). Hence other yield measures can only enter alternatively but not in addition to the interest rate. This is at odds with a bunch of money demand studies referring to Friedman (1956) where a whole bandwidth of asset returns determines money demand. Although such approaches have been successful in fitting the data, those empirical findings are “non-structural” (as Lucas (1980) formulates it) from the perspective of the standard CCAPM framework.

---

3 The safe interest rate can of course be substituted by a more refined measure of the opportunity costs of money. Note that the income elasticity can differ from one. Equalizing \(\sigma\) and \(\gamma\) implies the standard unit income elasticity. Moreover, for low interest rate levels, \(\ln\left(\frac{i_t}{1 + i_t}\right)\) can be approximated by \(i_t\) which leads to the standard semi-log specification often used in empirical applications.

4 This point, though in a less fierce manner, was already made by Friedman (1956) who reasoned that due to arbitrage among different yields, interest rates should serve as a sufficient statistic to describe opportunity costs of holding money.
2.2 Uncertainty

Empirical work based on the latter framework has to quantify the evolution of the variable $b_t$ in (2.11). Following the motivation of our study we assume that this liquidity shift factor is driven by macroeconomic uncertainty. However, in the economic profession there is no universally valid definition of “uncertainty”. Rather, the concept of uncertainty may be one with the most numerous and diverse definitions attached to it. In economic theory, uncertainty is usually associated with the stochastic environment that agents live in. For instance, in the theory of choice, uncertainty occurs in a lottery which assigns certain payoffs to certain states of the world that are not yet known to the decision maker. For example, in a CAPM-world the uncertainty associated with a certain asset may be measured by the conditional second moments of excess return and consumption. Frequently, this notion of uncertainty is used synonymously to risk.

However, this is not the type of uncertainty that is relevant in our money demand specification. According to the model presented in the preceding section, this type of risk should enter into the first order condition (2.5) as it influences the expected covariance between the stochastic discount factor and the return of a risky asset. In other words, in this general model framework, risk should ceteris paribus be solely incorporated in the price of risky assets. As a consequence, since $i_t$ is sufficient to describe the evolution of money there should be no direct link between risk and money holdings. Such a link can only exist indirectly if the degree of risk influences the development of the safe interest rate or output. Thus, risk in this sense should not carry extra information about money holdings.

Thus, turning back to the initial motivation of our study, uncertainty should be conceived as a possibly broader concept. Reflecting the interpretation of the ECB, we treat uncertainty as the bundle of forces that have contributed to a shift in preference for liquidity. These include high suffered stock market losses, high experienced volatility and unsettling geopolitical events. These experiences are supposed to have disturbed investors general confidence and led to a general preference for less risky investments in bonds and money. The meaning of uncertainty as it is used in our context may be closely linked to “pessimism” and the German word “Verunsicherung”. The latter denotes the result of being made feeling insecure, possibly due to bad experiences in the past.

Against this background, in the literature, such measures of uncertainty are usually constructed from backward looking measures such as variances over moving windows. Choi and Oh (2003) and Attah-Mensah (2004), for instance, assume that $b_t$ is a linear combination of variables such as experienced volatility of GDP, money, exchange rates
and others. Whereas, in our study we mainly concentrate on financial market measures which in the discussion were supposed to be linked to the specific characteristics of recent developments. Those are volatility and low returns on stock markets, the correlation between bond and stock returns, corporate bonds spreads, but also sentiment indicators derived from surveys.

3 Empirical Methods

Against the background of the preceding discussion based on equation (2.11), we develop an empirical model explaining real money holdings taking into account the influence of measures of macroeconomic uncertainty. Therefore, we first construct a small dynamic model with unobserved components which extracts measures of uncertainty from a set of variables representing financial market developments and economic sentiment data. In a second step, the extracted measures serve as additional covariables in a money demand error-correction model.

3.1 Empirical Model and Estimation Approach

We specify the long-run money demand equation as

\[(m - p)_{t} - a_0 - a_1 y_t - a_2 oc_t - b \cdot \tilde{unc}_{1t} = u_t, \quad (3.12)\]

where \(m - p\) is log real money balances, \(y\) is log real income, and \(oc\) is an opportunity cost variable. The \(I(1)\) variable \(\tilde{unc}_{1t}\) represents economic uncertainty whereas the tilde on top indicates that this variable is an estimate stemming from our unobserved components model.

Short-run dynamics are captured by a vector-error-correction model. Let

\[z_t = ((m - p)_t, y_t, oc_t, \tilde{unc}_{1t})',\]

then

\[\Delta z_t = c_0 + \sum_{j=1}^{l} C_j \Delta z_{t-j} + c_1 EC_{t-1} + d \cdot \tilde{unc}_{2t} + \epsilon_t \quad (3.13)\]

with \(EC_{t-1} = u_{t-1}\) being the error correction term. Equation (3.13) allows for the possibility of introducing an additional \(I(0)\) uncertainty variable \(\tilde{unc}_{2t}\) to enter the short-run dynamics. The money demand system will be estimated by full-information maximum likelihood based on the Johansen procedure.
For the general specification of the model for the latent uncertainty measure, denote by \( x_t \) an \( n \times 1 \) vector of observable measurement variables. Each element of \( x \) is driven by a \( p < n \) dimensional vector of common factors \( unc \) and an idiosyncratic component \( w \). The elements of \( unc \) are assumed to be unobservable. For the \( i \)th element of \( x_t \) we have

\[
x_{it} = \sum_{j=1}^{p} b_{ij} unc_{jt} + w_{it}, \tag{3.14}
\]

or in matrix notation\(^5\)

\[
x_t = B unc_t + w_t, \tag{3.15}
\]

where \( B \) is an \( n \times p \) matrix. The \( i, j \)-element of \( B \) is referred to as the loading of factor \( j \) on variable \( i \). Some of the elements of \( x_t \) will be \( I(1) \), so at least one of the factors or the corresponding idiosyncratic component has to be \( I(1) \) as well.\(^6\)

For the evolution of the latent uncertainty process we assume a VAR(1)-process

\[
unc_t = \mathcal{K} unc_{t-1} + \eta_t. \tag{3.16}
\]

Factor innovations are independent and identically distributed as

\[
\eta_t \sim N(0,Q). \tag{3.17}
\]

Factors are assumed to be independent, hence \( \mathcal{K} \) and \( Q \) are diagonal.

Each of the idiosyncratic components evolves as an AR(1)-process

\[
w_{it} = \gamma_i w_{it-1} + \xi_{it}, \tag{3.18}
\]

where the \( \xi_{it} \) are each Gaussian white noise independent of \( \eta_t \),

\[
\xi_{it} \sim N(0, s_t^2). \tag{3.19}
\]

Equations (3.15) - (3.19) fully specify the general interaction between the unobservable common factors \( unc_t \) and the observable quantities \( x_t \).

For the estimation of the unobserved components model, it is cast into a state space form.\(^7\) The state vector consists of the common factors and the idiosyncratic components. The transition equation is given by

\[
\begin{pmatrix}
unc_t \\
w_t
\end{pmatrix} =
\begin{pmatrix}
\mathcal{K} & 0_{p \times n} \\
0_{n \times p} & \Gamma
\end{pmatrix}
\begin{pmatrix}
unc_{t-1} \\
w_{t-1}
\end{pmatrix}
+ 
\begin{pmatrix}
\eta_t \\
\xi_t
\end{pmatrix}, \tag{3.20}
\]

\(^5\) For the model that we use for estimation, the dimension of the vector \( x_t \) will actually change over time. This is due to missing observations for some of the measurements in \( x_t \) in the first quarters of our sample. We abstract from this here to keep the notation simple.

\(^6\) We use the terms “factor” and “latent uncertainty measure” interchangeably.

\(^7\) See Hamilton (1994).
with
\[
\begin{pmatrix}
\eta_t \\
\xi_t
\end{pmatrix} \sim N\left(0, \begin{pmatrix} Q & 0 \\ 0 & S \end{pmatrix}\right),
\]
(3.21)
where \(\xi_t = (\xi_{1t}, \ldots, \xi_{nt})'\), \(\Gamma = \text{diag}(\gamma_1, \ldots, \gamma_n)\), and \(S = \text{diag}(s_{1}^2, \ldots, s_{n}^2)\). In compact notation,
\[
\alpha_t = T \alpha_{t-1} + v_t, \quad v_t \sim N(0, V).
\]
(3.22)

The measurement equation relates observable variables to the state vector. If all \(n\) elements of \(x_t\) were observed at all times, the measurement equation would be given by
\[
x_t = (B|I_n) \alpha_t,
\]
(3.23)
where \(I_n\) is the \(n \times n\) identity matrix.\(^8\) However, for our euro area data set described below, the last two of the \(n = 6\) variables in \(x_t\) – the measures of consumer and industry confidence – are observed from time \(t^* > 1\) on only.\(^9\) Thus, the dimension of the measurement vector changes over time. It is given by \(x_t = (x_{1t}, \ldots, x_{4t})'\) for \(t = 1, \ldots, t^*\), and by \(x_t = (x_{1t}, \ldots, x_{6t})'\) for \(t = t^* + 1, \ldots, T\). Accordingly, the matrix relating the state and the measurement vector also changes over time. Hence, we write
\[
x_t = (B_t|I_{n_t}) \alpha_t,
\]
(3.24)
where for \(t = 1, \ldots, t^*\) the matrices \(B_t\) and \(I_{n_t}\) consist of the first four rows of \(B\) and \(I_n\), respectively. For \(t = t^* + 1, \ldots, T\), \((B_t|I_{n_t}) = (B|I_n)\).

Since our model is a linear Gaussian state space model, estimation of the unobservable paths of the latent factors can be conducted using the Kalman filter and maximum likelihood. Denote by \(X_s\) the sequence of observed measurement vectors until time \(s\), augmented by a vector \(x_0\) of constants, \(X'_s = (x_0, x_1, \ldots, x_s)\). The sequences of conditional densities \(\{p(\alpha_t|X'_t)\}\) (filtering densities), \(\{p(\alpha_t|X'_t-1)\}\) and \(\{p(x_t|X'_t-1)\}\) (prediction densities) are Gaussian and thus fully described by their means and variance-covariance matrices. Given a set of model parameters, say \(\psi\), these moments can be computed by the Kalman filter. Based on a sequence \(\{p(x_t|X'_t-1)\}\) of one-step prediction densities of the measurement vector the log-likelihood is given by
\[
l(\psi; X'_t) = \sum_{t=1}^{T} \ln p(x_t|X'_t-1),
\]
(3.25)
\(^8\)Note that the measurement equation contains no measurement error since the idiosyncratic components are part of the state vector.
\(^9\)That is, \(t^*\) corresponds to 1985Q1, see the data description below.
where the vector $\psi$ collects all unknown parameters. Maximising with respect to $\psi$ yields the ML estimator $\hat{\psi}$.\(^{10}\)

Kalman filtering based on the ML estimates $\hat{\psi}$ yields a sequence of conditional expectations $E(\alpha_t|X_t;\hat{\psi}) =: a_{tlt}$, $t = 1, \ldots, T$, as estimates of the state vectors. The first two components of $a_{tlt}$ constitute the estimated factor vector, which will be denoted by $\hat{w}_t$. The remaining components of $a_{tlt}$ are the estimated idiosyncratic components, say $\hat{w}_t$, which are of no further interest for the problem under consideration.

In order to identify the model parameters, a priori restrictions have to be imposed. First, it should be noted that there are no intercepts, neither in the measurement nor in the transition equation. This is justified since all variables in $x_t$ have their mean subtracted. If intercepts are included in the measurement equation, their estimates turn out to be statistically insignificant.

Second, since some of the variables in $x_t$ can be stationary and some not, we can use this pattern for our identification scheme. Because of the structure of (3.15) these properties must transfer to the factor processes. If there are nonstationary measurements it follows that at least one of the factors has to be nonstationary as well. Consequently, in the case of $p = 2$ factors, we restrict the first autoregressive parameter of $K$ in (3.20) to be unity, and estimate the second

$$
K = \begin{pmatrix} 1 & 0 \\ 0 & \kappa_2 \end{pmatrix}.
$$

Third, related to that we assume that some of the factor loadings in the matrix $B$ are zero. Stationary measurements are only allowed to load on the stationary factors, nonstationary measurements on both. As an example, for $n = 6$ and $p = 2$, we restrict $B$ to have the following shape,

$$
B = \begin{pmatrix} b_{11} & b_{12} \\ \vdots & \vdots \\ b_{41} & b_{42} \\ 0 & b_{52} \\ 0 & b_{62} \end{pmatrix}.
$$

Fourth, the diagonal covariance matrix $Q$ of factor innovations in (3.17) is assumed to be the identity matrix. To justify this assumption, note that for a model with an arbitrary diagonal matrix $Q = \text{diag}(q_1^2, \ldots, q_n^2)$, factors can be rescaled by dividing by the

\(^{10}\)Due to the changing dimension of our measurement vector, the Kalman filter algorithm has to be suitably modified.
respective $q$. Appropriate scaling of the matrix $B$ leads to an observationally equivalent model, with $Q = I_p$.

Fifth, we assume that the variances of idiosyncratic components will be the same for all variables in $x_t$, that is $s^2_i = s^2$ for $i = 1, \ldots, n$ in (3.19). This is not an innocuous assumption per se. However, it is reasonable as our measurement variables are all normalised to have the same variance and the estimates of the autoregressive parameters of the idiosyncratic components turn out to be similar.

For our model used in the following section with $n = 6$ measurement variables and two factors, this leaves us with 18 free parameters to be estimated: the autoregressive parameter $\kappa_2$ for the factor evolution, the 10 non-zero elements in the loading matrix $B$, the variance $s^2$ of the innovations for the idiosyncratic components and their autoregressive parameters $\gamma_1, \ldots, \gamma_n$.

4 Euro area

In this section we apply the empirical model developed in the preceding section to euro area data. First, we choose and describe the data for the money demand and the uncertainty extraction model. Second, the model is estimated.

4.1 Variables and Data

For the standard money demand variables we use quarterly euro area data from 1980Q1 to 2004Q4. Data for the M3 money stock, the GDP-deflator, real GDP and the three-month interest rate are taken from an updated version of the data base in Fagan, Henry, and Mestre (2001) (Area Wide Model) and official ECB statistics. Additionally, a series of the own rate of return of M3 was taken from Calza et al. (2001) and extended by own calculations based on Bundesbank data. Using this, a measure of opportunity costs of real M3 holdings was calculated as the difference between the three-month interest rate and the own rate. Finally, all variables have been converted to logs. As the literature is not unanimous about taking logs in the case of interest rates, we have considered two variants, $oc_t = r_t^a - r_t^m$ where $oc_t^a$ is the three-month interest rate $r_t^a$ and $r_t^m$ is the own rate of return on M3, and alternatively $oc_t = \ln(r_t^a - r_t^m)$. Note that in the first case the coefficient on $oc_t$ in (3.12) can be interpreted as a semi-elasticity, whereas in the second case it represents the elasticity of money with respect to $r_t^a - r_t^m$. Figure 1 shows the four time series of $(m - p)_t$, $yt$, $r_t^a - r_t^m$ and $\ln(r_t^a - r_t^m)$. 


Figure 1: Log of M3, log of real income, opportunity cost in levels, and opportunity cost in logs.

For constructing the measurement variables to be used in the unobserved components model, we build on raw data of monthly or daily frequency which are finally converted to quarterly frequency by computing 3-month averages. Those raw data are transformed to six measures: a measure of the correlation between bond and stock returns, $SBCORR$, a measure of losses experienced on European stock market, $SLOSS$, a measure for stock market volatility, $SVOLA$, a measure of stock market returns $SRET$ and measures of consumer and industry confidence, $CCONF$ and $ICONF$. Together those indicators form the vector of observable variables in equation (3.15), i.e.

$$
x_t = \begin{pmatrix}
SBCORR_t \\
SLOSS_t \\
SVOLA_t \\
SRET_t \\
CCONF_t \\
ICONF_t
\end{pmatrix}.
$$

The variables in $x_t$ are shown in figures 2 and 3 whereas all variables are normalised by subtracting their respective sample mean and dividing by their sample standard deviation.

The correlation between returns on long-term government bonds and stocks $SBCORR$
Figure 2: Correlation between stock and bond returns, stock market loss, stock market volatility and stock market return.

can be viewed as a measure of time-varying risk perception. A negative correlation points to a period of relatively higher risk perception as market participants tend to substitute less risky bonds for stocks in their portfolios. Periods of moderate risk perception, in contrast, tend to be characterised by a positive correlation between stock and bond returns, as a decrease in interest rates is likely to spur stock prices while increasing bond returns at the same time.\textsuperscript{11} Hence alternatively, this indicator can simply be viewed as a direct measure representing shifts from bonds to stocks which certainly could be induced by high uncertainty. The measure $SBCORR$ is constructed by computing the correlation between month-to-month returns on stocks and bonds over a rolling window of 36 months. Bond returns are computed using the German bond price performance index REX, stock returns are based on a euro area stock market performance index.\textsuperscript{12}

The second measure $SLOSS$ represents experienced stock market losses over a medium-term horizon. The time series is constructed as a moving window of suffered losses over

\textsuperscript{11}See the box entitled “Risk aversion and developments in monetary aggregates” in ECB (2004) and the references cited therein.

\textsuperscript{12}The bond index is from the Bundesbank data base. The stock market index used here and in the following is based on a series provided at a daily frequency by Datastream.
a period of about three years. To be specific, first day-to-day returns are computed using the daily stock market index. Each day \( t \) is then assigned the five percent quantile of the distribution of the last 780 returns, i.e. the highest losses over that period.\(^{13}\)

For capturing stock market volatility, the variable \( SVOLA \) is a robust variance measure of daily stock market returns over a rolling window of 780 days. The variance measure is taken as the inter-quartile range, i.e. the difference between the 75 percent and the 25 percent quantile of returns within the respective window. The variable \( SRET \) is a medium-term stock market return. We compute it as the 12-quarter change in the logarithm of quarterly values of our euro area stock market index.

Finally, the measures \( CCONF \) and \( ICONF \), are the consumer and industry confidence indicators constructed by the European Commission.\(^{14}\) Note that these two time series are available from 1985Q1 on, only.

Following the discussion in 3.1 for the specification of the uncertainty model the order of integration of the measurement variables has to be assessed. Unit root tests showed that the first four series constructed in this section are clearly I(1) while the sentiment indices should rather be treated as I(0).\(^{15}\)

\(^{13}\)Since 5 days per week data are used, 260 days roughly comprise a year, thus, 780 days represent a three-year horizon.


\(^{15}\)Unit root tests based on ADF, Phillips-Perron and KPSS, not reported.
4.2 Empirical Results

The following shows the parameter estimates for the unobserved components model which was specified to contain two unobservable factors. Standard errors are given in parentheses. According to the the unit root properties of the data the last two entries in the first column have been set to zero a priori,

\[
\hat{B} = \begin{pmatrix}
-0.18 & 0.084 \\
(0.025) & (0.030) \\
0.064 & 0.0088 \\
(0.019) & (0.019) \\
0.051 & 0.0045 \\
(0.018) & (0.019) \\
-0.19 & -0.069 \\
(0.025) & (0.029) \\
0.0 & -0.29 \\
(-) & (0.031) \\
0.0 & -0.34 \\
(-) & (0.033)
\end{pmatrix}, \quad \hat{K} = \begin{pmatrix}
1 & 0 \\
0 & 0.928 \\
\end{pmatrix}, \quad \hat{s} = 0.17.
\]

The first factor loads significantly on all of the first four measurement variables. The second factor exhibits significant loadings for the measure of stock returns, the measure of correlation between stock and bond returns and for the measures of industry and consumer confidence.

As regards the AR-process in (3.20) and (3.22) which drives the uncertainty factor, the first diagonal element \( K_{11} \) is restricted to be equal to unity as the first factor is assumed to be I(1). The element \( K_{22} \) is estimated as 0.928, hence it is very close to one which implies a high persistence of this second uncertainty factor.

Figure 4 contains the paths of the estimated uncertainty measures, \( unc_{1t} \) and \( unc_{2t} \). It stands out that the first measure exhibits a sharp rise towards the end of the sample. This mainly reflects the decrease in the stock returns and bond stock return correlation as well as the increase in losses and volatility displayed in figure 2. The second measure, which is primarily associated with the confidence indicators, shows its highest peak in 1993. This reflects the euro area wide recession at that time initiated by the cooling down after the German reunification boom and accompanied by the turbulences within the European Exchange Rate Mechanism. Moreover, the graphs demonstrate the imposed unit root properties of the two factors. The first rather resembles a nonstationary process while the second seems to be I(0).

The estimated factors are now employed to augment the money demand equation. The first factor which is I(1) enters the long-run money demand relationship. For the error correction model we try two variants: one, in which the second factor shows up in
the short-run dynamics \((d \neq 0\) in (3.13)), and another in which it does not \((d = 0\). For
the opportunity cost variable we employ both, the level and the log-specification, i.e.,

\[
\text{oc}_t = r_t^s - r_t^m \quad \text{and} \quad \text{oc}_t = \ln(r_t^s - r_t^m). \tag{4.27}
\]

To get normally distributed residuals, we introduce dummy variables into the short-
run specification. For the two variants with opportunity costs in levels, we have a dummy
variable which is one for 1981Q2 and zero elsewhere, while for the two other specifications
we have a dummy variable for 1998Q4. Results of the estimation of the four specifications
are provided in table 1.

We first consider specifications 1 and 3. Including our uncertainty measure establishes
a cointegration relation between real money, income, opportunity costs and the uncertainty
proxy. This holds true for both the level- and the log-specification of opportunity
costs. The parameters for income, opportunity costs and the uncertainty measure all have
the expected sign and are significantly (at the five percent level) different from zero. At
the five-percent level, the lambda trace test suggests one cointegration relation. The error
correction term significantly enters the dynamics of \(\Delta(m - p)_t\). Uncorrelatedness and
normality of the residuals cannot be rejected.\(^{16}\)

\(^{16}\)The p-values of the LM test for serial correlation may appear somewhat small. However, looking at
the graphs of the (cross-)autocorrelation functions (not reported here) there is no indication of severe
misspecification.
**Table 1: M3 money demand, euro area, 1980Q1 - 2004Q4**

<table>
<thead>
<tr>
<th>Specification</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
</tr>
</thead>
<tbody>
<tr>
<td>( oc_t )</td>
<td>level</td>
<td>level</td>
<td>log</td>
<td>log</td>
</tr>
<tr>
<td>( unc_{2t} ) in ECM</td>
<td>no</td>
<td>yes</td>
<td>no</td>
<td>yes</td>
</tr>
</tbody>
</table>

| Lag length \( l \) in ECM | 2 | 2 | 2 | 2 |

### Long-run relationship

| \( y_t \) | 1.10 (0.067) | 1.26 (0.052) | 1.20 (0.058) | 1.30 (0.043) |
| \( oc_t \) | -2.43 (0.43) | -1.20 (0.34) | -0.062 (0.014) | -0.032 (0.010) |
| \( unc_{1t} \) | 0.68 (0.11) | 0.71 (0.085) | 0.39 (0.11) | 0.61 (0.077) |

### \( \lambda \) trace statistic\(^\dagger\)

| \( r = 0 \) | 49.0 | 58.0 | 42.2 | 51.5 |
| 0.05 crit. val. | 41.1 | 41.1 |
| \( r \leq 1 \) | 19.2 | 21.8 | 19.1 | 13.8 |
| 0.05 crit. val. | 23.8 | 23.8 |

### Error correction, selected results

| \( EC_{t-1} \) | -0.093 (0.025) | -0.15 (0.027) | -0.085 (0.029) | -0.15 (0.028) |
| \( unc_{2t} \) | - | 0.088 (0.028) | - | 0.098 (0.030) |
| LM(1) | 0.16 | 0.09 | 0.06 | 0.15 |
| LM(4) | 0.13 | 0.20 | 0.08 | 0.09 |
| Joint Jarque-Bera | 0.16 | 0.25 | 0.12 | 0.20 |
| \( R^2 \Delta(m - p)_t \) | 0.28 | 0.40 | 0.25 | 0.37 |
| \( \overline{R}^2 \Delta(m - p)_t \) | 0.20 | 0.32 | 0.16 | 0.29 |

\(^\dagger\) Critical values are computed via simulation using DISCO 1.4 by Bent Nielsen. Note that these are valid for the cases without \( unc_{2t} \) entering the short-run specification.
Specifications 2 and 4 are identical to specifications 1 and 3, respectively, except that the second factor enters the short-run dynamics of the error correction model. Again, estimation is carried out using full-information maximum likelihood. The second factor apparently helps to explain the short-run dynamics of real money. The adjusted $R^2$ for the $\Delta(m - p)_t$ part of the ECM increases and the Schwarz criterion decreases. Normality and absence of autocorrelation of residuals cannot be rejected. Moreover, introducing the second factor into the short-run specification leads to an increase of the long-run income elasticity and a decrease (in absolute value) of the interest rate elasticity. This holds for both specifications of the opportunity cost variable. However, if one compares the cointegration residuals of specifications 1 and 2 as well as 3 and 4 in figures 5 and 6, they look very similar to each other. This suggests that incorporating an additional short-run factor in simultaneous estimation does not affect the long-run estimates substantially.

Finally, we compare the evolution of excess liquidity implied by the standard money demand specifications with those augmented by the uncertainty measures. Therefore, we estimate the standard model $((m - p)_t = a_0 + a_1y_t + a_2oc_t + u_t)$ without uncertainty using data from 1980Q1 to 2001Q4 by maximum likelihood. The sample is reduced as estimation over the whole data range leads to a break down of cointegration properties. The parameter estimates are then used to compute projected residuals for the rest of the sample range available.

Figures 5 and 6 show the comparison of cointegration residuals ($EC_{t-1}$) for two interest rate specifications. While figure 5 is based on the opportunity cost specification in levels, figure 6 uses the logarithmic version. Until 2001 the standard specification and the augmented specification give rise to similar residuals. From then on, however, the (unstable) standard specification indicates an increasing excess liquidity exceeding 7 percent since 2003Q2. This holds true for both specifications of the opportunity cost variable. The augmented specification, in contrast, does not exhibit such a rise in excess liquidity.

---

17 The asymptotic critical values for the $\lambda$-trace tests on cointegration in this setting will generally depend on the nature of the $unc_{2t}$ process. Thus, to this end we interpret specifications 2 and 4 as an augmentation of the short-run dynamics of a model, for which the existence of an equilibrium between $m - p$, $y$, $oc$ and $unc_{1t}$ has already been established.

18 See Carstensen (2004) for a formal test of this.
Figure 5: Excess liquidity ($EC_{t-1}$) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in levels.

Figure 6: Excess liquidity ($EC_{t-1}$) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in logs.
Thus, if in addition to income and opportunity cost an uncertainty measure is associated with real money balances, the extended model implies a higher demand for money in a period of increased uncertainty. This in turn leads to the conclusion that the amount of liquidity held by euro area residents in particular over the period 2001 to 2004 was roughly in line with long-run money demand and should be expected to decrease if - ceteris paribus - macroeconomic uncertainty departs from the scene.

5 The Case of the US

The preceding analysis shows that the exceptional increase in euro area money holdings can be related to high macroeconomic uncertainty in recent years. The potential problem of our model and its conclusion so far is that it is based on end-of-sample evidence. With the euro area data available we cannot test whether the supposed effect occurs systematically as there is no similar constellation elsewhere in the sample. Hence, in order to check for the robustness of our approach, we apply it to US data.

5.1 Monetary Measures for the US

Similar to the euro area, monetary growth in the US has been unusually strong in recent years. However, in contrast to the euro area, the existence of a stable money demand function had been questioned even longer ago. There is common evidence that such a relationship between money, output and opportunity costs had been prevailing only until the early nineties.\(^{19}\) For the following years, US money stock seemed to be too low in comparison to the historical relation with output and opportunity costs (“missing money”).\(^{20}\) There are several not necessarily competing explanations for this deviation. The main argument is that due to a process of financial innovations new asset types became available which served as attractive alternatives to other subaggregates of broad money. As those new assets carried higher yields in comparison to money, portfolio shifts away from old money into the new liquid assets occurred to a greater extent. This was particularly true for mutual funds especially for those based on bonds. These funds provided the opportunity to invest in relatively more risky longer-term assets with higher yields and to write checks on these investments at the same time.

\(^{19}\)See, e.g., Calza and Sousa (2003) for an overview. For the US, mostly M2 is used as the standard measure for broad money instead of M3 as in the case of the euro area.

\(^{20}\)The original notion of “missing money” was of course intended to characterise a similar situation in the late seventies. See Goldfeld and Sichel (1990).
Against this background, the observed structural break in US money demand functions was rather a discontinuity in the measurement of money than a fundamental economic shift. This interpretation follows from the results of Carlson, Hoffman, Keen, and Rasche (2000). They show that the demand for alternative broad money measures like M2M and MZM, which try to consider those shifts between some mutual funds and components of M2, has been stable even if the nineties are included in the period of analysis.\(^{21}\)

An alternative or additional interpretation follows from Choi and Oh (2003) who show that the structural change in the demand for US M1 can well be explained by measures of macroeconomic volatility and financial innovation. This holds for the marked missing money periods in the seventies and nineties. If this finding was also valid for broader monetary aggregates it would imply that significant shifts from money into other non-monetary assets occurred systematically. Hence, it would mean a more fundamental economic effect instead of the measurement problem discussed in the preceding paragraph. Moreover, a potential companion phenomenon implied by this study could be that the evolution of financial innovations which is supposed to cause a drain from money was fostered by macroeconomic uncertainty and risk effects. This complicates the analysis as not only the investor’s choice between money and alternatives would become endogenous but also the creation of those alternative assets.

Against the background of this discussion we first test whether our uncertainty approach could explain the behavior of the standard broad monetary aggregate M2. This resembles the approach of Choi and Oh (2003) and hence could point at a systematically missing influence in standard money demand functions for broad US monetary aggregates, in particular for the missing money periods. Second, we apply our approach to the demand for monetary aggregates used in Carlson et al. (2000) that try to take into account the presumed shift among liquid assets in the nineties.

### 5.2 The US Uncertainty Model

For the application to the US, similar measurement variables as for the euro area are being used.\(^{22}\) To be specific, the measurement vector comprises 6 variables, a stock

\(^{21}\)In contrast to M2, in the definition of M2M and MZM small time deposits are exempt. Additionally, MZM comprises institutional money market mutual funds.

\(^{22}\)Data in this section for the US are generally taken from the database of the Federal Reserve of St. Louis (FRED). Stock data are from Datastream. The consumer confidence index is provided by the University of Michigan. All computations are based on monthly data which are converted into quarterly observations by taking averages.
return volatility measure\textsuperscript{23}, a loss measure\textsuperscript{24}, the correlation between stock and bond returns\textsuperscript{25} and the 3-year real stock return\textsuperscript{26}. Additionally, another standard measure related to economic uncertainty, the spread between corporate and government bond yields was introduced.\textsuperscript{27} Finally, a survey-based consumer sentiment index was employed representing public expectations.\textsuperscript{28} Data range from 1975Q1 to 2004Q4 which corresponds to the availability of monetary aggregates used later. Thus, the measurement vector has the following form,

\[
\begin{pmatrix}
SVOLA_t \\
SLOSS_t \\
SBCORR_t \\
SRET_t \\
SPREAD_t \\
CCONS_t
\end{pmatrix}
\]

The measure of stock return volatility in figure 7 exhibits three characteristic periods. The first is the marked increase after 1975 which is a consequence of the stock market decline following the first oil price shock in late 1973. The second is the trough in the middle of the nineties. Finally, the high level around the New Economy boom is visible. Interestingly, with this measure designed to be robust to outliers there is no exceptional increase observable around the crash in October 1987. The bond-stock-market correlation shows a marked fall to an unprecedented level after 2000 which seems to have found its bottom in 2002/2003. Three-year real stock returns in 1975 still suffered from the decline in 1973/74 with a recovery following in the late seventies. The sharp but isolated peak in 1987 supports the view that the stock market turbulences in fall 1987 were rather a short-term correction of overpriced stocks than a crash which exhibits influences on middle to longer-term developments. This parallels the assessment expressed by the volatility variable. Over the last ten years, the marked stock market up- and downswing related to the New Economy boom is clearly visible. Peculiar for the development of the spread

\textsuperscript{23}The volatility measure at time $t$ is computed as the variance of preceding stock returns over a horizon of 3 years based on the S&P 500 index. The variance measure is chosen to be the difference between the 75\% and 25\% quantile in order to provide an outlier robust measure.

\textsuperscript{24}It is computed as a moving window of the lower 5\% quantile of the above defined stock returns over the respective last three years.

\textsuperscript{25}Computed as the correlation between those returns over a horizon of three years. The bond return is based on a bond price performance index.

\textsuperscript{26}Real returns are based on the GDP-deflator.

\textsuperscript{27}Corporate bond yields are corresponding to Moody’s AAA.

\textsuperscript{28}The minor differences in the selection of measurement variables in comparison to the euro area are due to data availability. However, this turned out not to be crucial for the overall results.
between corporate AAA and long-term government bonds yields is the period 2000 to 2003
with a remarkably high level. The consumer sentiment indicator shows troughs around

As regards unit root properties, all series except for the consumer sentiment index are
clearly I(1). The consumer confidence index is a border case which is at least very close
to being I(1). As the assessment of the degree of integration is crucial for our identification
scheme, we follow a two-pronged approach. In the first case, we treat consumer sentiment
as I(0) and use the identification pattern with one I(1)- and one I(0)-factor employed
above for the euro area model. Second, the sentiment index is assumed to be I(1). In the

---

20Order of integration results are based on ADF and KPSS tests, not reported.
latter case we opt for the extraction of only one uncertainty factor which is consequently specified as being I(1).

The general specification of the US unobserved components model is equal to the one used in section 3.1. Hence, we estimate the model 3.15 and 3.16 with a serially correlated idiosyncratic components \( w_{it} \),

\[
x_t = B \text{unc}_t + w_t, \\
\text{unc}_t = \mathcal{K} \text{unc}_{t-1} + \eta_t.
\]

Applying the two factor identification pattern, we obtain the following estimates of the \( B \) and \( \mathcal{K} \)-matrix:\(^{30}\)

\[
\hat{B} = \begin{pmatrix}
0.13 & -0.1 \\
0.12 & -0.18 \\
-0.14 & 0.04 \\
-0.08 & 0.18 \\
0.13 & -0.06 \\
-0.17 & \\
\end{pmatrix}, \quad \hat{\mathcal{K}} = \begin{pmatrix}
1 & 0 \\
0 & 0.97 \\
\end{pmatrix}, \quad \hat{s}^2 = 0.1.
\]

Similar to the first uncertainty factor for the euro area in figure 4, \( \text{unc}_{1t} \) in figure 8 shows a remarkable increase starting in 2001. As regards the second uncertainty factor \( \text{unc}_{2t} \) the estimate of 0.97 for the AR-parameter in \( \hat{\mathcal{K}} \) mirrors the unclear unit root property of the consumer sentiment indicator. This point estimate is not distinguishable from one which makes the second factor \( \text{unc}_{2t} \) look like an I(1)-variable. Moreover, the impression of non-stationarity is supported by inspection of figure 9.

In the second case with only one factor, the following estimates have been received,

\[
\hat{B} = \begin{pmatrix}
0.24 \\
0.24 \\
-0.13 \\
-0.11 \\
0.11 \\
-0.034 \\
\end{pmatrix}, \quad \mathcal{K} = 1, \quad \hat{s}^2 = 0.11.
\]

\(^{30}\)In the case of the US the AR parameters \( \gamma_i \) of the error process \( w_t \) was restricted to a maximum value of 0.75.
Figure 8: Uncertainty factors, two-factor model

Here, the loading of the uncertainty factor on the consumer sentiment variable is not significant. This might point at that an identification scheme with two factors independent of the unit root properties is more appropriate. However, we proceed with this version of the single factor $unc_t$ as leaving out consumer sentiment from the measurement vector does not change the other loadings or the shape of the extracted factor.\textsuperscript{31} Note that the first factor $unc_t$ from the 2-factor scheme looks very similar to the single factor from the pure I(1)-model. The only difference is that in the second case uncertainty seems to be more pronounced in the first quarters of the sample.

Figure 9: Uncertainty factor, one-factor model

\textsuperscript{31}Moreover, the money demand estimations in the following did not change qualitatively when using a single factor based on only 5 measurements.
5.3 US Money Demand Functions

For the US the semi-log version of the money demand model (3.12) and (3.13) is used,

$$\Delta(m - p)_t = \epsilon_0 + \sum_{j=1}^{l} C_j \Delta z_{t-j} + c_1 EC_{t-1} + \epsilon_t,$$

with

$$EC_t = (m - p)_t - a_0 - a_1 y_t - a_2 \delta c_t - b \cdot \vec{w} \epsilon c_t.$$

It relates money holdings \(m_t\) deflated by the GDP-deflator \(p_t\) to real GDP \(y_t\) and opportunity costs \(\delta c_t\) measured by the nominal three month interest rate minus the own rate of the respective monetary aggregate.\(^{32}\) Accordingly, the standard money demand function is augmented by the uncertainty measures estimated in the previous section.

Estimation and testing has been carried out in two steps. First, the long-run money demand was estimated by DOLS.\(^{33}\) Afterwards, we tested for cointegration by using augmented Dickey-Fuller unit root tests on the residuals. In the second step, the error correction model was estimated with OLS. The specifications vary in terms of estimation range and inclusion of the uncertainty measures.

The reason for not using the Johansen-methodology here is that its application leads to some econometric problems. Though deflated M2M and MZM generally exhibit a clear trend with output and opportunity costs in the long run, they are subject to significant short-run variations in some periods. This finally causes the residuals of the system’s estimation to be plagued by non-normality and serial correlation.\(^{34}\) Though evidence for cointegration is found, these problems are quite likely to bias Johansen-based estimates in small samples. As a consequence, it seems not reasonable to apply the Johansen ML-estimator for the employed US data. Hence, results based on the more robust Dynamic Ordinary Least Squares (DOLS) estimator by Saikkonen (1991) are reported.

Table 2 contains DOLS estimates of the long-run demand for real M2 together with unit roots tests for cointegration. Three specifications are displayed. The first relating real M2 holdings to output and opportunity costs shows that cointegration could be found over the period of 1975 to 1989. This date approximately marks the beginning of the significant growth of mutual funds and hence represents the break in M2 money demand.

\(^{32}\)Unit root tests not reported here reveal that all variables are I(1).

\(^{33}\)Estimates based on the fully modified estimator of Phillips and Hansen (1990) (FMOLS) were similar and hence not reported.

\(^{34}\)A simple cure, e.g. augmenting the system with a number of manageable and adequate dummies or removing outliers, does not lead to satisfactory results.
Extending the estimation period to 1998Q4 which is the sample end corresponding to Carlson et al. (2000) (revealing stable demand functions for M2M and MZM) leads to a break down of cointegration even if the M2 money demand function is augmented by an uncertainty factor. Finally, using the full sample and hence incorporating the period with extraordinary levels of our uncertainty variables does not change this finding.

Table 2: M2 money demand
first observation 1975Q1

<table>
<thead>
<tr>
<th>end of sample</th>
<th>1989Q4</th>
<th>2004Q4</th>
<th>1998Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_t$</td>
<td>0.891</td>
<td>0.640</td>
<td>0.623</td>
</tr>
<tr>
<td>(0.011)</td>
<td>(0.051)</td>
<td>(0.055)</td>
<td></td>
</tr>
<tr>
<td>$oc_t$</td>
<td>-2.600</td>
<td>-3.387</td>
<td>-3.649</td>
</tr>
<tr>
<td>(0.189)</td>
<td>(1.214)</td>
<td>(1.401)</td>
<td></td>
</tr>
<tr>
<td>$unc_{1t}$</td>
<td>-</td>
<td>0.659</td>
<td>0.493</td>
</tr>
<tr>
<td>(0.261)</td>
<td>(0.595)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>unit root</td>
<td>-4.55</td>
<td>-3.00</td>
<td>-3.05</td>
</tr>
</tbody>
</table>

standard errors in parentheses, ADF unit root test, critical values (5%) of unit root tests for cointegration: -3.74 (k=3), -4.1 (k=4), k = number of variables in cointegration relation

Overall, the results in table 2 clearly show that augmenting a standard M2 money demand specification by our uncertainty measure does not restore the cointegration property between money, output and opportunity costs existent before the nineties. Hence, assuming we provide a correct measure, we would reject the possibility that macroeconomic uncertainty had caused a shift away from liquid to other non or less liquid assets in the nineties. Comparing with Carlson et al. (2000), the results are rather in line with the view that a shift among liquid assets took place.

Tables 3 and 4 contain the results of the estimation of the money demand model for the monetary aggregate M2M using the different factors extracted before. For all estimations we chose different sample ends: the last quarter of 2004 (full sample), autumn 2001 which coincides with the terror attacks in New York and hence is one of the possible dates marking the beginning of a phase with increased uncertainty and end of 1998, which corresponds to the date used in Carlson et al. (2000).

---

35These estimates are based on the first factor $unc_{1t}$ of the 2-factor uncertainty model. Using the other I(1)-uncertainty measure from the one-factor specification provided similar results as regards cointegration properties.

36For expository reasons in the following regressions the estimated uncertainty factors depicted in figures 8 and 9 were divided by 100 in order to receive higher numerical values for the point estimates.
The first estimates in table 3 compare the standard money demand specification with the augmented one where the factor $unc_{1t}$ from our first uncertainty model is used. Over the whole sample and the one ending in 2001Q3 we cannot reject the hypothesis of no cointegration between real M2M, GDP and M2M-opportunity costs. Similar to Carlson et al. (2000) we find a cointegrated long-run relation for the sample ranging until 1998Q4 where signs and magnitude of the parameters are within the expected range. Augmenting the long-run money demand function by the uncertainty factor $unc_{1t}$ leads to restoring the cointegration property for both of the extended samples. Not unexpectedly, the relevant uncertainty measure $unc_{1t}$ is not distinguishable from zero for the shortest sample while it is significant and positive for the sample until 2001Q3. This could either imply that the impact of uncertainty was not relevant before 1998 or that this measure rather works as a threshold variable which only has an impact when a certain level is exceeded.

Table 3: M2M money demand, 2 factors - factor 1 only
first observation 1975Q1

<table>
<thead>
<tr>
<th>end of sample</th>
<th>2004Q4</th>
<th>2001Q3</th>
<th>1998Q4</th>
<th>2004Q4</th>
<th>2001Q3</th>
<th>1998Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_t$</td>
<td>0.996</td>
<td>0.938</td>
<td>0.864</td>
<td>0.949</td>
<td>0.923</td>
<td>0.877</td>
</tr>
<tr>
<td></td>
<td>(0.097)</td>
<td>(0.050)</td>
<td>(0.022)</td>
<td>(0.033)</td>
<td>(0.026)</td>
<td>(0.019)</td>
</tr>
<tr>
<td>$oc_{1t}$</td>
<td>-4.652</td>
<td>-4.379</td>
<td>-4.547</td>
<td>-3.591</td>
<td>-3.905</td>
<td>-4.457</td>
</tr>
<tr>
<td></td>
<td>(1.416)</td>
<td>(0.700)</td>
<td>(0.271)</td>
<td>(0.409)</td>
<td>(0.321)</td>
<td>(0.235)</td>
</tr>
<tr>
<td>$unc_{1t}$</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>0.988</td>
<td>0.652</td>
<td>0.000</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(0.144)</td>
<td>(0.157)</td>
<td>(0.165)</td>
</tr>
<tr>
<td>unit root</td>
<td>-2.27</td>
<td>-3.15</td>
<td>-5.26</td>
<td>-4.68</td>
<td>-5.17</td>
<td>-6.65</td>
</tr>
</tbody>
</table>

| lags          | 2      | 2      | 2      | 2      | 2      | 2      |
| $EC_{t-1}$    | -0.086 | -0.163 | -0.408 | -0.255 | -0.383 | -0.674 |
|              | (0.033)| (0.050)| (0.074)| (0.067)| (0.082)| (0.094)|
| $R^2$         | 0.53   | 0.55   | 0.63   | 0.58   | 0.61   | 0.70   |

standard errors in parentheses, ADF unit root test, critical values (5\%) of unit root tests for cointegration: -3.74 (k=3), -4.1 (k=4), k = number of variables in cointegration relation

Conducting the same exercise with either both factors $unc_{1t}$ and $unc_{2t}$ together or with $unc_{2t}$ only, unambiguously showed that the second uncertainty measure is not significant and moreover does not help to reestablish cointegrated long-run relations over the extended samples.\textsuperscript{37} Thus, $unc_{1t}$ seems to be the relevant measure of uncertainty while

\textsuperscript{37}Results not shown here are available upon request.
unc₂ₜ is of minor importance.³⁸

Using the second specification of the uncertainty model which extracts only one I(1)-factor and augmenting money demand leads to the results in table 4. They corroborate our findings that the uncertainty measure restores cointegration in comparison to the standard version. Moreover, the point estimates do not differ much from the ones in table 3. This also demonstrates that our uncertainty extraction procedure is relatively robust.³⁹

<table>
<thead>
<tr>
<th>Table 4: M2M money demand, 1 factor</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>first observation 1975Q1</td>
<td></td>
</tr>
<tr>
<td>end of sample</td>
<td>2004Q4</td>
</tr>
<tr>
<td>long-run relation</td>
<td></td>
</tr>
<tr>
<td>yt</td>
<td>1.046</td>
</tr>
<tr>
<td></td>
<td>(0.030)</td>
</tr>
<tr>
<td>ocₜ</td>
<td>-4.046</td>
</tr>
<tr>
<td></td>
<td>(0.383)</td>
</tr>
<tr>
<td>unc₁₂ₜ</td>
<td>1.039</td>
</tr>
<tr>
<td></td>
<td>(0.160)</td>
</tr>
<tr>
<td>unit root</td>
<td>-4.56</td>
</tr>
<tr>
<td>ECM</td>
<td></td>
</tr>
<tr>
<td>lags</td>
<td>2</td>
</tr>
<tr>
<td>ECₜ₋₁</td>
<td>-0.237</td>
</tr>
<tr>
<td></td>
<td>(0.069)</td>
</tr>
<tr>
<td>R²</td>
<td>0.56</td>
</tr>
</tbody>
</table>

standard errors in parentheses, ADF unit root test, critical values (5%) of unit root tests for cointegration: -3.74 (k=3), -4.1 (k=4), k = number of variables in cointegration relation

Finally, in table 5 the same regressions are carried out with MZM as a measure for broad money. As the alternative approaches lead to very similar results, we only report those for the version where the variable unc₁₂ₜ from the 2-factor uncertainty model is used to augment long-run money demand. Though the results are similar to the ones for M2M they are a little bit less supportive for our augmented money demand approach at first sight. In comparison to the M2M case the general level of the unit root test values is a bit higher, i.e. cointegration is less likely to be found. For the full sample, the unit

³⁸The factor unc₂ₜ might have some impact on short term monetary developments as its difference appears to be significant in the short-run dynamics in the ECM-model. Certainly, the empirical non-relevance may also be related to the unclear time series properties mentioned above.

³⁹The results are also very similar if the consumer sentiment is dropped from the list of measurements and money demand is then augmented by using the respective factor from a 5 variable uncertainty model.
root test statistic for cointegration within the augmented model marginally misses the 5% critical value. However, cointegration can be established at a level of 10% (critical value = -3.81).\textsuperscript{40} Overall, we notwithstanding interpret these results as additional support for the view that the recent development of uncertainty at least contributed to the increase in US liquidity. This assessment is also encouraged by the following graphs.

Table 5: MZM money demand, 2 factors, factor 1 only

<table>
<thead>
<tr>
<th>end of sample</th>
<th>2004Q4</th>
<th>2001Q3</th>
<th>1998Q4</th>
<th>2004Q4</th>
<th>2001Q3</th>
<th>1998Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>( y_t )</td>
<td>1.226</td>
<td>1.135</td>
<td>1.023</td>
<td>1.144</td>
<td>1.114</td>
<td>1.031</td>
</tr>
<tr>
<td>( (0.200) )</td>
<td>( (0.122) )</td>
<td>( (0.030) )</td>
<td>( (0.053) )</td>
<td>( (0.048) )</td>
<td>( (0.024) )</td>
<td></td>
</tr>
<tr>
<td>( (2.953) )</td>
<td>( (1.714) )</td>
<td>( (0.371) )</td>
<td>( (0.664) )</td>
<td>( (0.595) )</td>
<td>( (0.363) )</td>
<td></td>
</tr>
<tr>
<td>( unc_{1t} )</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>1.552</td>
<td>1.037</td>
<td>-0.070</td>
</tr>
<tr>
<td>( (0.224) )</td>
<td>( (0.278) )</td>
<td>( (0.205) )</td>
<td>( (0.224) )</td>
<td>( (0.278) )</td>
<td>( (0.205) )</td>
<td></td>
</tr>
<tr>
<td>unit root</td>
<td>-1.68</td>
<td>-1.78</td>
<td>-4.31</td>
<td>-4.06</td>
<td>-3.80</td>
<td>-5.62</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>lags</th>
<th>2</th>
<th>2</th>
<th>2</th>
<th>2</th>
<th>2</th>
<th>2</th>
</tr>
</thead>
<tbody>
<tr>
<td>ECM</td>
<td>0.047</td>
<td>-0.072</td>
<td>-0.326</td>
<td>-0.130</td>
<td>-0.183</td>
<td>-0.573</td>
</tr>
<tr>
<td>( (0.023) )</td>
<td>( (0.039) )</td>
<td>( (0.069) )</td>
<td>( (0.051) )</td>
<td>( (0.067) )</td>
<td>( (0.094) )</td>
<td></td>
</tr>
<tr>
<td>( R^2 )</td>
<td>0.51</td>
<td>0.51</td>
<td>0.59</td>
<td>0.54</td>
<td>0.55</td>
<td>0.66</td>
</tr>
</tbody>
</table>

standard errors in parentheses, ADF unit root test, critical values (5%) of unit root tests for cointegration: -3.74 (k=3), -4.1 (k=4), k = number of variables in cointegration relation

To visualise the impact of augmenting US money demand functions by the uncertainty variable, figures 10 and 11 display the monetary overhang, i.e. the deviation of money holdings from its long-run equilibrium, using M2M and MZM respectively. The “standard” version is based on the estimation of the not-augmented specifications in table 3 and 5 with the sample ending in 1998. Values for the following periods are then projected using these estimates and the data for the rest of the sample. This is compared to the overhang based on the specification using the \( unc_{1t} \)-factor over the full data sample in the respective tables. It is apparent from the figures that including the uncertainty variable largely helps

\textsuperscript{40}The fact that it is a bit more difficult to reveal a cointegration relation for MZM than M2M can possibly be explained by the relatively higher volatility in the dynamics of MZM in comparison to M2M. This possibly influences the DOLS estimates as the embedded correction for serial correlation might work less perfect. Thus the unit root tests might perform “less well” in favour of our hypothesised cointegration relation.
to explain the excess money holdings that would follow from the view of 1998. Similar to the results for the euro area, this comparison suggests that the high degree of uncertainty fostered a shift from risky to liquid assets in the US in recent years.

Figure 10: Monetary overhang, M2M

Figure 11: Monetary overhang, MZM
6 Conclusions

In this paper we show that a measure of macroeconomic uncertainty largely explains the increase in holdings of euro area M3 over the period from 2001 to 2004. Therefore, uncertainty indicators with different unit root properties are extracted from a set of variables describing financial market characteristics and economic sentiment. Augmenting a standard money demand model by such an obtained I(1)-indicator establishes a cointegration relationship for samples that include recent periods. Additionally, a second stationary factor, mainly representing consumer and industry confidence, significantly improves the empirical description of the short-run growth of real M3. Moreover, it is demonstrated that similar mechanisms seem to be at work in the US. An application of the approach to US data enhances the explanation of recent developments of the broad monetary aggregates M2M and MZM.

By deriving uncertainty factors the study provides a new empirical proxy representing the impact of liquidity preference shifts on euro area M3. Accordingly, the main developments reflecting macroeconomic uncertainty and hence driving monetary growth in recent years were connected with low returns, high average volatility and losses on stock markets. Thus, the empirical model captures and supports the portfolio shift interpretation of the ECB of EMU monetary developments.
References


The following Discussion Papers have been published since 2004:

**Series 1: Economic Studies**

1. 2004 Foreign Bank Entry into Emerging Economies: An Empirical Assessment of the Determinants and Risks Predicated on German FDI Data
   Torsten Wezel

   Torsten Wezel

3. 2004 Policy Instrument Choice and Non-Coordinated Monetary Policy in Interdependent Economies
   Giovanni Lombardo
   Alan Sutherland

4. 2004 Inflation Targeting Rules and Welfare in an Asymmetric Currency Area
   Giovanni Lombardo

5. 2004 FDI versus cross-border financial services: The globalisation of German banks
   Claudia M. Buch
   Alexander Lipponer

6. 2004 Clustering or competition? The foreign investment behaviour of German banks
   Claudia M. Buch
   Alexander Lipponer

7. 2004 PPP: a Disaggregated View
   Christoph Fischer

8. 2004 A rental-equivalence index for owner-occupied housing in West Germany 1985 to 1998
   Claudia Kurz
   Johannes Hoffmann

9. 2004 The Inventory Cycle of the German Economy
   Thomas A. Knetsch

10. 2004 Evaluating the German Inventory Cycle Using Data from the Ifo Business Survey
    Thomas A. Knetsch

11. 2004 Real-time data and business cycle analysis in Germany
    Jörg Döpke
<table>
<thead>
<tr>
<th>Page</th>
<th>Year</th>
<th>Title</th>
<th>Authors</th>
</tr>
</thead>
<tbody>
<tr>
<td>12</td>
<td>2004</td>
<td>Business Cycle Transmission from the US to Germany – a Structural Factor Approach</td>
<td>Sandra Eickmeier</td>
</tr>
<tr>
<td>13</td>
<td>2004</td>
<td>Consumption Smoothing Across States and Time: International Insurance vs. Foreign Loans</td>
<td>George M. von Furstenberg</td>
</tr>
<tr>
<td>15</td>
<td>2004</td>
<td>Welfare Implications of the Design of a Currency Union in Case of Member Countries of Different Sizes and Output Persistence</td>
<td>Rainer Frey</td>
</tr>
<tr>
<td>16</td>
<td>2004</td>
<td>On the decision to go public: Evidence from privately-held firms</td>
<td>Ekkehart Boehmer, Alexander Ljungqvist</td>
</tr>
<tr>
<td>17</td>
<td>2004</td>
<td>Who do you trust while bubbles grow and blow? A comparative analysis of the explanatory power of accounting and patent information for the market values of German firms</td>
<td>Fred Ramb, Markus Reitzig</td>
</tr>
<tr>
<td>18</td>
<td>2004</td>
<td>The Economic Impact of Venture Capital</td>
<td>Astrid Romain, Bruno van Pottelsberghe</td>
</tr>
<tr>
<td>19</td>
<td>2004</td>
<td>The Determinants of Venture Capital: Additional Evidence</td>
<td>Astrid Romain, Bruno van Pottelsberghe</td>
</tr>
<tr>
<td>21</td>
<td>2004</td>
<td>How effective are automatic stabilisers? Theory and results for Germany and other OECD countries</td>
<td>Michael Scharnagl, Karl-Heinz Tödter</td>
</tr>
<tr>
<td>Page</td>
<td>Year</td>
<td>Title</td>
<td>Authors</td>
</tr>
<tr>
<td>------</td>
<td>------</td>
<td>-----------------------------------------------------------------------</td>
<td>------------------------------------------------------------------------</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Implications for the ECB</td>
<td></td>
</tr>
<tr>
<td>23</td>
<td>2004</td>
<td>Financial Liberalization and Business Cycles: The Experience of</td>
<td>Lúcio Vinhas, de Souza</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Countries in the Baltics and Central Eastern Europe</td>
<td></td>
</tr>
<tr>
<td>24</td>
<td>2004</td>
<td>Towards a Joint Characterization of Monetary Policy and the Dynamics</td>
<td>Ralf Fendel</td>
</tr>
<tr>
<td></td>
<td></td>
<td>of the Term Structure of Interest Rates</td>
<td></td>
</tr>
<tr>
<td>25</td>
<td>2004</td>
<td>How the Bundesbank really conducted monetary policy: An analysis</td>
<td>Christina Gerberding, Andreas Worms, Franz Seitz</td>
</tr>
<tr>
<td></td>
<td></td>
<td>based on real-time data</td>
<td></td>
</tr>
<tr>
<td>26</td>
<td>2004</td>
<td>Real-time Data for Norway: Challenges for Monetary Policy</td>
<td>T. Bernhardsen, Ø. Eitrheim, A.S. Jore, Ø. Røisland</td>
</tr>
<tr>
<td>27</td>
<td>2004</td>
<td>Do Consumer Confidence Indexes Help Forecast Consumer Spending in</td>
<td>Dean Croushore</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Real Time?</td>
<td></td>
</tr>
<tr>
<td>28</td>
<td>2004</td>
<td>The use of real time information in Phillips curve relationships for</td>
<td>Maritta Paloviita, David Mayes</td>
</tr>
<tr>
<td></td>
<td></td>
<td>the euro area</td>
<td></td>
</tr>
<tr>
<td>29</td>
<td>2004</td>
<td>The reliability of Canadian output gap estimates</td>
<td>Jean-Philippe Cayen, Simon van Norden</td>
</tr>
<tr>
<td>30</td>
<td>2004</td>
<td>Forecast quality and simple instrument rules - a real-time data</td>
<td>Heinz Glück, Stefan P. Schleicher</td>
</tr>
<tr>
<td></td>
<td></td>
<td>approach</td>
<td></td>
</tr>
<tr>
<td>31</td>
<td>2004</td>
<td>Measurement errors in GDP and forward-looking monetary policy: The</td>
<td>Peter Kugler, Thomas J. Jordan, Carlos Lenz, Marcel R. Savioz</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Swiss case</td>
<td></td>
</tr>
<tr>
<td>Year</td>
<td>Page</td>
<td>Title</td>
<td>Author(s)</td>
</tr>
<tr>
<td>------</td>
<td>------</td>
<td>----------------------------------------------------------------------</td>
<td>-----------------------------------</td>
</tr>
<tr>
<td>2004</td>
<td>32</td>
<td>Estimating Equilibrium Real Interest Rates in Real Time</td>
<td>Todd E. Clark</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Sharon Kozicki</td>
</tr>
<tr>
<td>2004</td>
<td>33</td>
<td>Interest rate reaction functions for the euro area</td>
<td>Karsten Ruth</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Evidence from panel data analysis</td>
<td></td>
</tr>
<tr>
<td>2004</td>
<td>34</td>
<td>The Contribution of Rapid Financial Development to Asymmetric Growth of Manufacturing Industries: Common Claims vs. Evidence for Poland</td>
<td>George M. von Furstenberg</td>
</tr>
<tr>
<td>2004</td>
<td>35</td>
<td>Fiscal rules and monetary policy in a dynamic stochastic general equilibrium model</td>
<td>Jana Kremer</td>
</tr>
<tr>
<td>2004</td>
<td>36</td>
<td>Inflation and core money growth in the euro area</td>
<td>Manfred J.M. Neumann</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Claus Greiber</td>
</tr>
<tr>
<td>2004</td>
<td>37</td>
<td>Taylor rules for the euro area: the issue of real-time data</td>
<td>Dieter Gerdesmeier</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Barbara Roffia</td>
</tr>
<tr>
<td>2004</td>
<td>38</td>
<td>What do deficits tell us about debt?</td>
<td>Jürgen von Hagen</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Empirical evidence on creative accounting with fiscal rules in the EU</td>
<td>Guntram B. Wolff</td>
</tr>
<tr>
<td>2004</td>
<td>39</td>
<td>Optimal lender of last resort policy in different financial systems</td>
<td>Falko Fecht</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Marcel Tyrell</td>
</tr>
<tr>
<td>2004</td>
<td>40</td>
<td>Expected budget deficits and interest rate swap spreads - Evidence for France, Germany and Italy</td>
<td>Kirsten Heppke-Falk</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Felix Hüfner</td>
</tr>
<tr>
<td>2004</td>
<td>41</td>
<td>Testing for business cycle asymmetries based on autoregressions with a Markov-switching intercept</td>
<td>Malte Knüppel</td>
</tr>
<tr>
<td>2005</td>
<td>1</td>
<td>Financial constraints and capacity adjustment in the United Kingdom – Evidence from a large panel of survey data</td>
<td>Ulf von Kalckreuth</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>Emma Murphy</td>
</tr>
<tr>
<td>Page</td>
<td>Year</td>
<td>Title</td>
<td>Authors</td>
</tr>
<tr>
<td>------</td>
<td>------</td>
<td>----------------------------------------------------------------------</td>
<td>------------------------------------------------------------------------</td>
</tr>
<tr>
<td>2</td>
<td>2005</td>
<td>Common stationary and non-stationary factors in the euro area analyzed in a large-scale factor model</td>
<td>Sandra Eickmeier</td>
</tr>
<tr>
<td>3</td>
<td>2005</td>
<td>Financial intermediaries, markets, and growth</td>
<td>F. Fecht, K. Huang, A. Martin</td>
</tr>
<tr>
<td>4</td>
<td>2005</td>
<td>The New Keynesian Phillips Curve in Europe: does it fit or does it fail?</td>
<td>Peter Tillmann</td>
</tr>
<tr>
<td>5</td>
<td>2005</td>
<td>Taxes and the financial structure of German inward FDI</td>
<td>Fred Ramb, A. J. Weichenried</td>
</tr>
<tr>
<td>6</td>
<td>2005</td>
<td>International diversification at home and abroad</td>
<td>Fang Cai, Francis E. Warnock</td>
</tr>
<tr>
<td>7</td>
<td>2005</td>
<td>Multinational enterprises, international trade, and productivity growth: Firm-level evidence from the United States</td>
<td>Wolfgang Keller, Steven R. Yeaple</td>
</tr>
<tr>
<td>8</td>
<td>2005</td>
<td>Location choice and employment decisions: a comparison of German and Swedish multinationals</td>
<td>S. O. Becker, K. Ekholm, R. Jäckle, M.-A. Muendler</td>
</tr>
<tr>
<td>9</td>
<td>2005</td>
<td>Business cycles and FDI: evidence from German sectoral data</td>
<td>Claudia M. Buch, Alexander Lipponer</td>
</tr>
<tr>
<td>10</td>
<td>2005</td>
<td>Multinational firms, exclusivity, and the degree of backward linkages</td>
<td>Ping Lin, Kamal Saggi</td>
</tr>
<tr>
<td>11</td>
<td>2005</td>
<td>Firm-level evidence on international stock market comovement</td>
<td>Robin Brooks, Marco Del Negro</td>
</tr>
<tr>
<td>12</td>
<td>2005</td>
<td>The determinants of intra-firm trade: in search for export-import magnification effects</td>
<td>Peter Egger, Michael Pfaffermayr</td>
</tr>
</tbody>
</table>
13 2005 Foreign direct investment, spillovers and absorptive capacity: evidence from quantile regressions Sourafel Girma Holger Görg

14 2005 Learning on the quick and cheap: gains from trade through imported expertise James R. Markusen Thomas F. Rutherford

15 2005 Discriminatory auctions with seller discretion: evidence from German treasury auctions Jörg Rocholl

16 2005 Consumption, wealth and business cycles: why is Germany different? B. Hamburg, M. Hoffmann, J. Keller

17 2005 Tax incentives and the location of FDI: evidence from a panel of German multinationals Thiess Buettner Martin Ruf

18 2005 Monetary Disequilibria and the Euro/Dollar Exchange Rate Dieter Nautz Karsten Ruth

19 2005 Berechnung trendbereinigter Indikatoren für Deutschland mit Hilfe von Filterverfahren Stefan Stamfort

20 2005 How synchronized are central and east European economies with the euro area? Sandra Eickmeier Jörg Breitung

21 2005 Asymptotic distribution of linear unbiased estimators in the presence of heavy-tailed stochastic regressors and residuals J.-R. Kurz-Kim S.T. Rachev G. Samorodnitsky

22 2005 The Role of Contracting Schemes for the Welfare Costs of Nominal Rigidities over the Business Cycle Matthias Pastian

<table>
<thead>
<tr>
<th>Year</th>
<th>Title</th>
<th>Author</th>
</tr>
</thead>
<tbody>
<tr>
<td>2005</td>
<td>Forecasting German GDP using alternative factor models based on large datasets</td>
<td>Christian Schumacher</td>
</tr>
<tr>
<td>2005</td>
<td>Time-dependent or state-dependent price setting? – micro-evidence from German metal-working industries –</td>
<td>Harald Stahl</td>
</tr>
<tr>
<td>2005</td>
<td>Money demand and macroeconomic uncertainty</td>
<td>Claus Greiber</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Wolfgang Lemke</td>
</tr>
</tbody>
</table>
### Series 2: Banking and Financial Studies

<table>
<thead>
<tr>
<th>Year</th>
<th>Title</th>
<th>Authors</th>
</tr>
</thead>
<tbody>
<tr>
<td>2004</td>
<td>Forecasting Credit Portfolio Risk</td>
<td>A. Hamerle, T. Liebig, H. Scheule</td>
</tr>
<tr>
<td></td>
<td>Systematic Risk in Recovery Rates – An Empirical Analysis of US Corporate Credit Exposures</td>
<td>Klaus Düllmann, Monika Trapp</td>
</tr>
<tr>
<td>2004</td>
<td>Does capital regulation matter for bank behaviour? Evidence for German savings banks</td>
<td>Frank Heid, Daniel Porath, Stéphanie Stolz</td>
</tr>
<tr>
<td></td>
<td>German bank lending during emerging market crises: A bank level analysis</td>
<td>F. Heid, T. Nestmann, B. Weder di Mauro, N. von Westernhagen</td>
</tr>
<tr>
<td>2004</td>
<td>How will Basel II affect bank lending to emerging markets? An analysis based on German bank level data</td>
<td>T. Liebig, D. Porath, B. Weder di Mauro, M. Wedow</td>
</tr>
<tr>
<td>2004</td>
<td>Estimating probabilities of default for German savings banks and credit cooperatives</td>
<td>Daniel Porath</td>
</tr>
<tr>
<td>2005</td>
<td>Measurement matters – Input price proxies and bank efficiency in Germany</td>
<td>Michael Koetter</td>
</tr>
<tr>
<td>2005</td>
<td>The supervisor’s portfolio: the market price risk of German banks from 2001 to 2003 – Analysis and models for risk aggregation</td>
<td>Christoph Memmel, Carsten Wehn</td>
</tr>
<tr>
<td>2005</td>
<td>Do banks diversify loan portfolios? A tentative answer based on individual bank loan portfolios</td>
<td>Andreas Kamp, Andreas Pfingsten, Daniel Porath</td>
</tr>
<tr>
<td>2005</td>
<td>Banks, markets, and efficiency</td>
<td>F. Fecht, A. Martin</td>
</tr>
<tr>
<td>Page</td>
<td>Year</td>
<td>Title</td>
</tr>
<tr>
<td>------</td>
<td>------</td>
<td>----------------------------------------------------------------------</td>
</tr>
<tr>
<td>5</td>
<td>2005</td>
<td>The forecast ability of risk-neutral densities of foreign exchange</td>
</tr>
<tr>
<td>6</td>
<td>2005</td>
<td>Cyclical implications of minimum capital requirements</td>
</tr>
<tr>
<td>7</td>
<td>2005</td>
<td>Banks’ regulatory capital buffer and the business cycle: evidence for German savings and cooperative banks</td>
</tr>
<tr>
<td>8</td>
<td>2005</td>
<td>German bank lending to industrial and non-industrial countries: driven by fundamentals or different treatment?</td>
</tr>
</tbody>
</table>
Visiting researcher at the Deutsche Bundesbank

The Deutsche Bundesbank in Frankfurt is looking for a visiting researcher. Visitors should prepare a research project during their stay at the Bundesbank. Candidates must hold a Ph D and be engaged in the field of either macroeconomics and monetary economics, financial markets or international economics. Proposed research projects should be from these fields. The visiting term will be from 3 to 6 months. Salary is commensurate with experience.

Applicants are requested to send a CV, copies of recent papers, letters of reference and a proposal for a research project to:

Deutsche Bundesbank
Personalabteilung
Wilhelm-Epstein-Str. 14

D - 60431 Frankfurt
GERMANY