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Money demand and macroeconomic uncertainty

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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

Während der letzen drei Jahre lag der Anstieg von M3 im Euro-Raum ständig über dem Referenzwert für das Geldmengenwachstum. Dadurch konnten bis dato stabile Geldnachfragefunktionen die monetäre Entwicklung nicht mehr erklären. Zudem offenbarten diese Funktionen Anzeichen von Instabilität. Eine Erklärung für dieses Phänomen könnte darin liegen, dass in einer Situation verstärkter makroökonomischer Unsicherheit in Verbindung mit niedrigen Erträgen auf Finanzaktiva die Liquiditätspräferenz erhöht ist. Unsicherheit ist jedoch im Hinblick auf eine ökonomische Analyse schwer zu erfassen, sowohl aus konzeptioneller als auch aus Sicht der Quantifizier- und Messbarkeit. Die meisten Ansätze in der Literatur verwenden einen bestimmten beobachtbaren Indikator als Näherungsgröße für diese Unsicherheit. Die Beschränkung auf eine bestimmte Variable erscheint jedoch teilweise etwas arbiträr. Im Gegensatz dazu wird in unserem Ansatz Unsicherheit als ein nicht beobachtbarer Prozess behandelt, der sich in mehreren messbaren ökonomischen Indikatoren widerspiegelt. Auf Basis solcher Maßzahlen schätzen wir dann die zu Grunde liegende Unsicherheit mit Hilfe eines Unbeobachtete Komponentenmodells. Diese beobachtbaren Indikatoren stellen im Wesentlichen Finanzmarktdaten dar, wie mittelfristige Erträge, Verluste und Volatilitäten. Es werden aber auch Umfragedaten über Industrie- und Konsumentenvertrauen verwendet. Die Analyse zeigt, dass die gewonnenen Unsicherheitsmaße helfen, den Anstieg von M3 im Euro-Raum in den Jahren 2001 bis 2004 zu erklären. Insbesondere kann eine kointegrierte Geldnachfragebeziehung für Stichproben, welche diesen Zeitraum berücksichtigen, nachgewiesen werden. Die Robustheit unseres Ansatzes wird mit Hilfe von US-Daten überprüft. Demnach schienen in den USA ähnliche Prozesse am Werke zu sein. Erweiterungen von Geldnachfragespezifikationen für bestimmte US-Geldmengenaggregate um Unsicherheitsmaße verbessert auch deren Erklärungsgehalt. Somit unterstützt diese Studie die Einschätzung, dass das monetäre Wachstum in den letzten Jahren durch Portfolioumschichtungen von riskanten in liquide Anlagen beeinflusst worden ist, welche durch eine erhöhte ökonomische Unsicherheit ausgelöst wurden.

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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 feeded 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."[1]

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.

[1]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

$$\max E_0 \left[\sum_{t=0}^{\infty} \beta^t u \left(C_t, \frac{M_t}{P_t} \right) \right] \tag{2.1}$$

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

$$C_t + q_t Q_t + \frac{B_t}{P_t} + \frac{M_t}{P_t} = (q_t + d_t) Q_{t-1} + R_{t-1} \frac{B_{t-1}}{P_t} + \frac{M_{t-1}}{P_t} + Y_t. \tag{2.2}$$

Defining the stochastic discount factor as

$$\mathcal{M}_{t+1} := \beta \frac{u_C(C_{t+1}, M_{t+1}/P_{t+1})}{u_C(C_t, M_t/P_t)}, \tag{2.3}$$

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

$$E_t \left[\mathcal{M}_{t+1} R_t \right] = 1, \tag{2.4}$$

$$E_t \left[\mathcal{M}_{t+1} \frac{q_{t+1} + d_{t+1}}{q_t} \right] = 1 \tag{2.5}$$

and

$$E_t \left[\mathcal{M}_{t+1} \cdot \frac{P_t}{P_{t+1}} \right] + \frac{u_{M/P}(C_t, M_t/P_t)}{u_C(C_t, M_t/P_t)} = 1. \tag{2.6}$$

Assuming a Fisher relationship

$$R_t = \frac{P_t}{P_{t+1}} (1 + i_t), \tag{2.7}$$

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

$$E_t \left[\mathcal{M}_{t+1} \frac{P_t}{P_{t+1}} \right] = \frac{1}{1 + i_t}. \tag{2.8}$$

[2] See Kim (2000), Ireland (2004), Attah-Mensah (2004), Baba (2000) and Petursson (2000).

Inserting into (2.5) and rearranging one obtains

$$\frac{u_{M/P}(C_t, M_t/P_t)}{u_C(C_t, M_t/P_t)} = \frac{i_t}{1+i_t}. \tag{2.9}$$

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

$$u(C_t, M_t/P_t) = \frac{C_t^{1-\sigma} - 1}{1-\sigma} + b_t^{\delta} \frac{(M_t/P_t)^{1-\gamma} - 1}{1-\gamma}, \tag{2.10}$$

where σ and γ represent coefficients of relative risk aversion with respect to consumption and real balances. The random process b_t is a preference shift variable with δ 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

$$\ln(M_t/P_t) = \frac{\sigma}{\gamma} \ln(C_t) - \frac{1}{\gamma} ln \left(\frac{i_t}{1+i_t} \right) + \frac{\delta}{\gamma} \ln(b_t). \tag{2.11}$$

This equation states that real balances are determined by consumption (being equal to output in equilibrium) and the safe interest rate.[3] 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.[4]

[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 σ and γ 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

$$(m-p)_t - a_0 - a_1 y_t - a_2 oc_t - b \cdot \widetilde{unc}_{1t} = u_t, \tag{3.12}$$

where $m-p$ is log real money balances, y is log real income, and oc is an opportunity cost variable. The I(1) variable $\widetilde{unc}_{1t}$ 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

$$z_t = ((m-p)_t, y_t, oc_t, \widetilde{unc}_{1t})',$$

then

$$\Delta z_t = c_0 + \sum_{j=1}^{l} C_j \Delta z_{t-j} + c_1 EC_{t-1} + d \cdot \widetilde{unc}_{2t} + \epsilon_t \tag{3.13}$$

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 $\widetilde{unc}_{2t}$ 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 ith 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

$$w_{it} = \gamma_i w_{i\,t-1} + \xi_{it}, \tag{3.18}$$

where the ξ_{it} are each Gaussian white noise independent of η_t,

$$\xi_{it} \sim N(0, s_i^2). \tag{3.19}$$

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

$$\begin{pmatrix} \eta_t \\ \xi_t \end{pmatrix} \sim N\left(0, \begin{pmatrix} Q & 0 \\ 0 & S \end{pmatrix}\right), \tag{3.21}$$

where $\xi_t = (\xi_{1t}, \ldots, \xi_{nt})'$, $\Gamma = diag(\gamma_1, \ldots, \gamma_n)$, and $S = diag(s_1^2, \ldots, s_n^2)$. In compact notation,

$$\alpha_t = \mathcal{T}\alpha_{t-1} + v_t, \quad v_t \sim N(0, V). \tag{3.22}$$

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

$$x_t = (B|I_n)\ \alpha_t, \tag{3.23}$$

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

$$x_t = (B_t|I_{n_t})\ \alpha_t, \tag{3.24}$$

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 $\mathcal{X}_s$ the sequence of observed measurement vectors until time s, augmented by a vector x_0 of constants, $\mathcal{X}_s = (x_0, x_1, \ldots, x_s)$. The sequences of conditional densities $\{p(\alpha_t|\mathcal{X}_t)\}$ (filtering densities), $\{p(\alpha_t|\mathcal{X}_{t-1})\}$ and $\{p(x_t|\mathcal{X}_{t-1})\}$ (prediction densities) are Gaussian and thus fully described by their means and variance-covariance matrices. Given a set of model parameters, say ψ^0, these moments can be computed by the Kalman filter. Based on a sequence $\{p(x_t|\mathcal{X}_{t-1})\}$ of one-step prediction densities of the measurement vector the log-likelihood is given by

$$l(\psi; \mathcal{X}_t) = \sum_{t=1}^{T} \ln p(x_t|\mathcal{X}_{t-1}), \tag{3.25}$$

[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 ψ collects all unknown parameters. Maximising with respect to ψ yields the ML estimator $\hat{\psi}$.[10]

Kalman filtering based on the ML estimates $\hat{\psi}$ yields a sequence of conditional expectations $E(\alpha_t|\mathcal{X}_t;\hat{\psi}) =: a_{t|t}$, $t = 1,\dots,T$, as estimates of the state vectors. The first two components of $a_{t|t}$ constitute the estimated factor vector, which will be denoted by $\widetilde{unc}_t$. The remaining components of $a_{t|t}$ are the estimated idiosyncratic components, say $\tilde{w}_t$, 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 $\mathcal{K}$ in (3.20) to be unity, and estimate the second

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

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,

$$B = \begin{pmatrix} b_{11} & b_{12} \\ \vdots & \vdots \\ b_{41} & b_{42} \\ 0 & b_{52} \\ 0 & b_{62} \end{pmatrix}. \tag{3.26}$$

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 $Q = diag(q_1^2,\dots,q_p^2)$, 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_i. 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 $s_i^2 = s^2$ 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 ideosyncratic 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 κ_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, $oc_t = r_t^s - r_t^m$ where oc_t^s is the three-month interest rate r_t^s and r_t^m is the own rate of return on M3, and alternatively $oc_t = \ln(r_t^s - r_t^m)$. 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 $r_t^s - r_t^m$. Figure 1 shows the four time series of $(m-p)_t$, y_t, $r_t^s - r_t^m$ and $\ln(r_t^s - r_t^m)$.

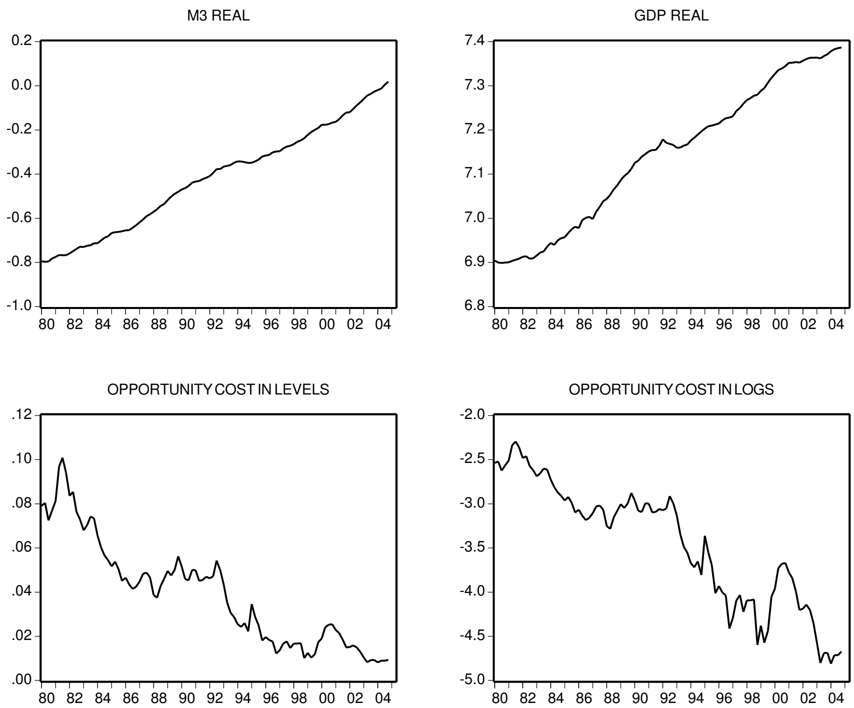

Figure 1: Log of M3, log of real income, opportunity cost in levels, and opportunity cost in logs.

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 $SBCORR$

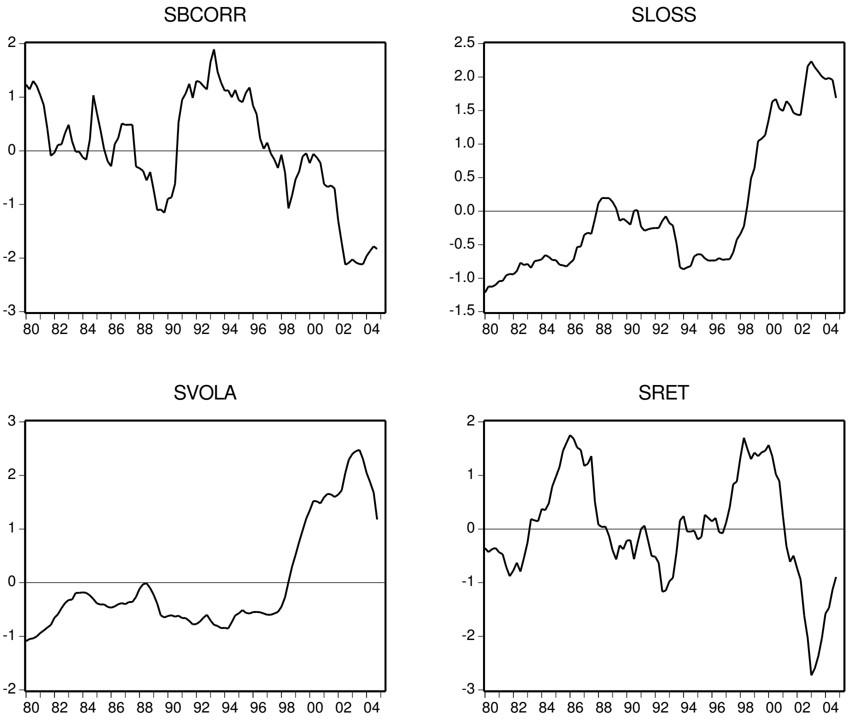

Figure 2: Correlation between stock and bond returns, stock market loss, stock market volatility and stock market return.

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.[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.[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

[11]See the box entitled "Risk aversion and developments in monetary aggregates" in ECB (2004) and the references cited therein.

[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.

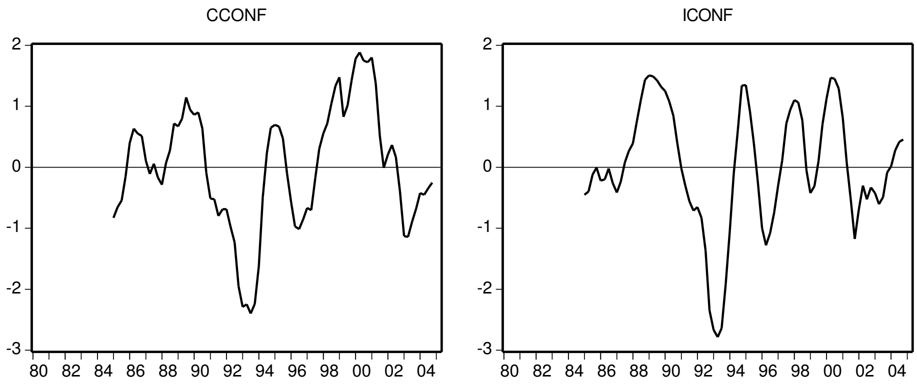

Figure 3: Consumer confidence and industry confidence.

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.

[14]See `http://europa.eu.int/comm/economy_finance/indicators/businessandconsumersurveys_en.htm`.

[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,

$$\hat{B} = \begin{pmatrix} \underset{(0.025)}{-0.18} & \underset{(0.030)}{0.084} \\ \underset{(0.019)}{0.064} & \underset{(0.019)}{0.0088} \\ \underset{(0.018)}{0.051} & \underset{(0.019)}{0.0045} \\ \underset{(0.025)}{-0.19} & \underset{(0.029)}{-0.069} \\ \underset{(-)}{0.0} & \underset{(0.031)}{-0.29} \\ \underset{(-)}{0.0} & \underset{(0.033)}{-0.34} \end{pmatrix}, \quad \hat{\mathcal{K}} = \begin{pmatrix} 1 & 0 \\ 0 & \underset{(0.038)}{0.928} \end{pmatrix}, \quad \hat{s} = \underset{(0.0063)}{0.17}.$$

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 $\mathcal{K}_{11}$ is restricted to be equal to unity as the first factor is assumed to be I(1). The element $\mathcal{K}_{22}$ 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

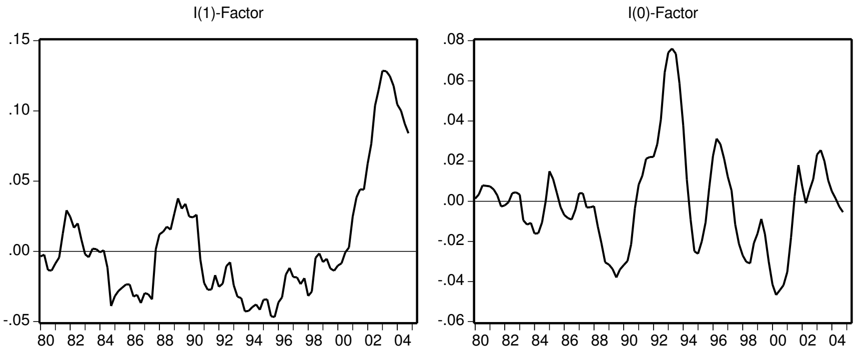

Figure 4: Estimated factors representing uncertainty

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.,

$$oc_t = r_t^s - r_t^m \quad \text{and} \quad oc_t = \ln(r_t^s - r_t^m). \tag{4.27}$$

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 cointegraton 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

Specification	1	2	3	4
oc_t	level	level	log	log
unc_{2t} in ECM	no	yes	no	yes
Lag length l in ECM	2	2	2	2
Dummy in ECM	1981Q2	1981Q2	1998Q4	1998Q4
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)
const.	-8.18	-9.39	-9.19	-9.80
λ trace statistic†				
$r=0$	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
$\bar{R}^2$ $\Delta(m-p)_t$	0.20	0.32	0.16	0.29
SC system	-19.54	-19.58	-21.63	-21.82

† 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.[17] 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 ($(m-p)_t = a_0 + a_1 y_t + a_2 oc_t + u_t$) 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.[18] 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 λ-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.

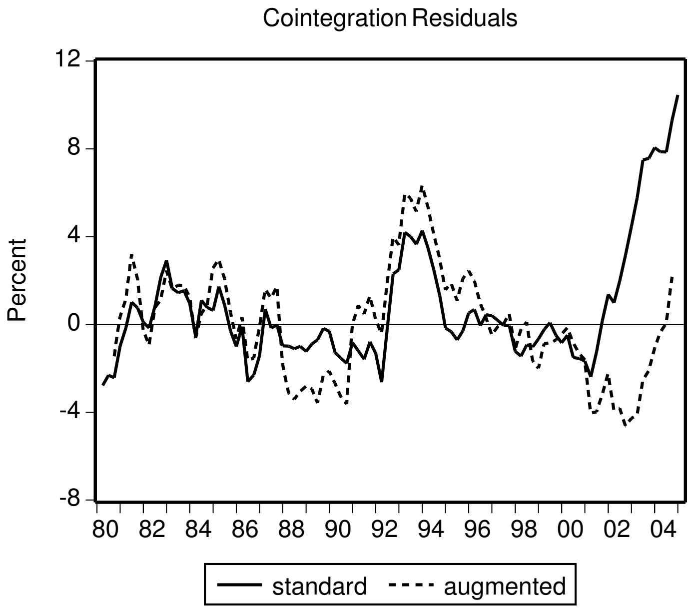

Figure 5: Excess liquidity (EC_{t-1}) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in levels.

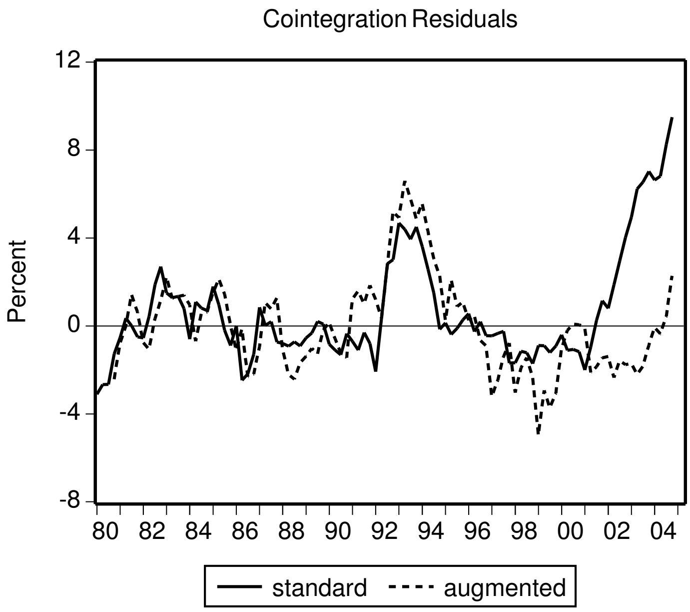

Figure 6: Excess liquidity (EC_{t-1}) implied by standard model and augmented model with uncertainty measures. Opportunity costs measured in logs.

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[23], a loss measure[24], the correlation between stock and bond returns[25] and the 3-year real stock return[26]. Additionally, another standard measure related to economic uncertainty, the spread between corporate and government bond yields was introduced.[27] Finally, a survey-based consumer sentiment index was employed representing public expectations.[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,

$$x_t = \begin{pmatrix} SVOLA_t \\ SLOSS_t \\ SBCORR_t \\ SRET_t \\ SPREAD_t \\ CCONS_t \end{pmatrix}.$$

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

[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.

[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.

[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.

[26] Real returns are based on the GDP-deflator.

[27] Corporate bond yields are corresponding to Moody's AAA.

[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.

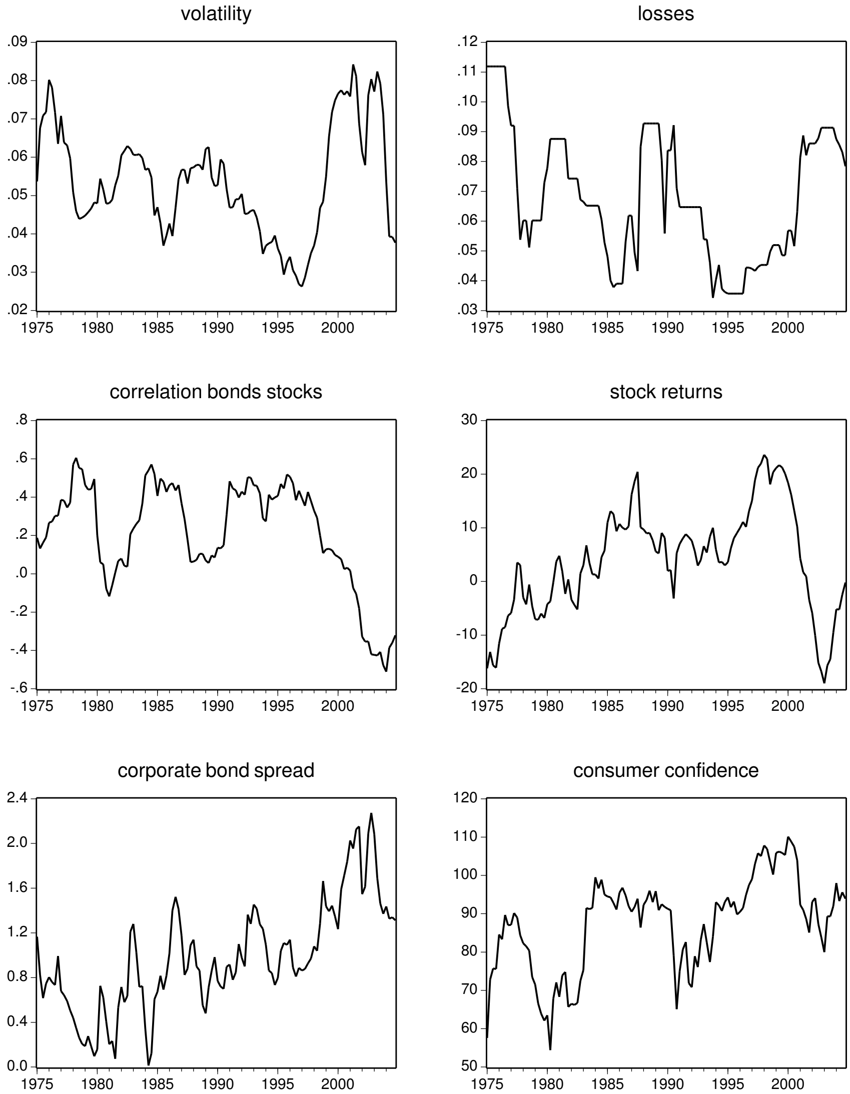

Figure 7: Measurements, US unobserved components model

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 1975, 1980 and 1991 while it hits "all time highs" around 2000.

As regards unit root properties, all series except for the consumer sentiment index are clearly I(1).[29] 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

[29] Order 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 ideosyncratic components w_{it},

$$\begin{aligned} x_t &= B\, unc_t + w_t, \\ unc_t &= \mathcal{K}\, unc_{t-1} + \eta_t. \end{aligned}$$

Applying the two factor identification pattern, we obtain the following estimates of the B and $\mathcal{K}$-matrix,[30]

$$\hat{B} = \begin{pmatrix} \underset{(0.033)}{0.13} & \underset{(0.039)}{-0.1} \\ \underset{(0.036)}{0.12} & \underset{(0.037)}{-0.18} \\ \underset{(0.024)}{-0.14} & \underset{(0.036)}{0.04} \\ \underset{(0.032)}{-0.08} & \underset{(0.032)}{0.18} \\ \underset{(0.031)}{0.13} & \underset{(0.042)}{-0.06} \\ - & \underset{(0.033)}{-0.17} \end{pmatrix}, \quad \hat{\mathcal{K}} = \begin{pmatrix} 1 & 0 \\ 0 & \underset{(0.022)}{0.97} \end{pmatrix}, \quad \hat{s}^2 = 0.1.$$

Similar to the first uncertainty factor for the euro area in figure 4, unc_{1t} in figure 8 shows a remarkable increase starting in 2001. As regards the second uncertainty factor 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 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,

$$\hat{B} = \begin{pmatrix} \underset{(0.035)}{0.24} \\ \underset{(0.035)}{0.24} \\ \underset{(0.029)}{-0.13} \\ \underset{(0.031)}{-0.11} \\ \underset{(0.034)}{0.11} \\ \underset{(0.035)}{-0.034} \end{pmatrix}, \quad \mathcal{K} = 1, \quad \hat{s}^2 = 0.11.$$

[30] In the case of the US the AR parameters γ_i of the error process w_t was restricted to a maximum value of 0.75.

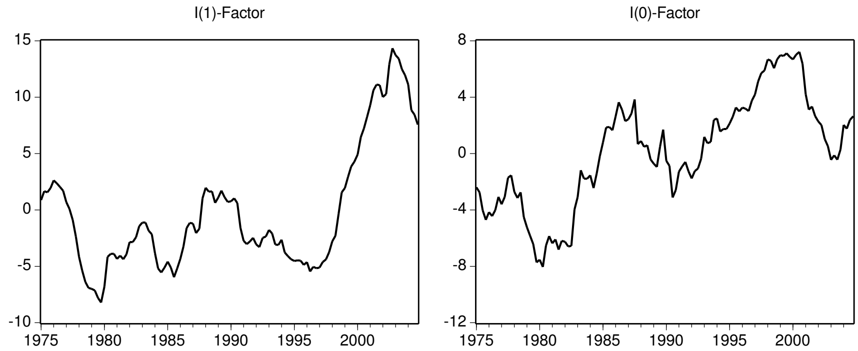

Figure 8: Uncertainty factors, two-factor model

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.[31] Note that the first factor unc_{1t} 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.

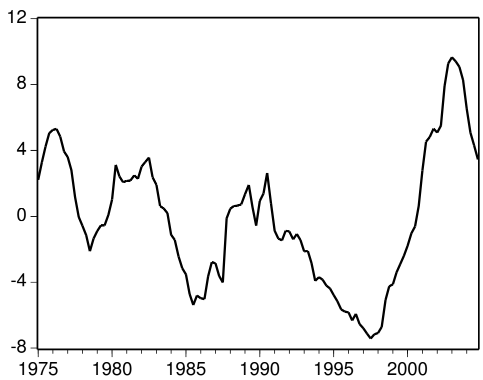

Figure 9: Uncertainty factor, one-factor model

[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,

$$\Delta(m-p)_t = c_0 + \sum_{j=1}^{l} C_j \Delta z_{t-j} + c_1 EC_{t-1} + \epsilon_t,$$

with

$$EC_t = (m-p)_t - a_0 - a_1 y_t - a_2 oc_t - b \cdot \widetilde{unc}_t.$$

It relates money holdings m_t deflated by the GDP-deflator p_t to real GDP y_t and opportunity costs oc_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.[35] 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

end of sample	1989Q4	2004Q4	1998Q4
	long-run relation		
y_t	0.891 (0.011)	0.640 (0.051)	0.623 (0.055)
oc_t	−2.600 (0.189)	−3.387 (1.214)	−3.649 (1.401)
unc_{1t}	-	0.659 (0.261)	0.493 (0.595)
unit root	−4.55	−3.00	−3.05

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.[36] 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).

[35] These 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.

[36] For expositional 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

end of sample	2004Q4	2001Q3	1998Q4	2004Q4	2001Q3	1998Q4
			long-run relation			
y_t	0.996 (0.097)	0.938 (0.050)	0.864 (0.022)	0.949 (0.033)	0.923 (0.026)	0.877 (0.019)
oc_t	-4.652 (1.416)	-4.379 (0.700)	-4.547 (0.271)	-3.591 (0.409)	-3.905 (0.321)	-4.457 (0.235)
unc_{1t}	-	-	-	0.988 (0.144)	0.652 (0.157)	0.000 (0.165)
unit root	-2.27	-3.15	-5.26	-4.68	-5.17	-6.65
			ECM			
lags	2	2	2	2	2	2
EC_{t-1}	-0.086 (0.033)	-0.163 (0.050)	-0.408 (0.074)	-0.255 (0.067)	-0.383 (0.082)	-0.674 (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.[37] Thus, unc_{1t} seems to be the relevant measure of uncertainty while

[37] Results not shown here are available upon request.

unc_{2t} 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 4: M2M money demand, 1 factor
first observation 1975Q1

end of sample	2004Q4	2001Q3	1998Q4
	long-run relation		
y_t	1.046 (0.030)	0.987 (0.034)	0.886 (0.029)
oc_t	−4.046 (0.383)	−4.178 (0.321)	−4.343 (0.204)
unc_{1t}	1.039 (0.160)	0.657 (0.218)	−0.001 (0.186)
unit root	−4.56	−4.72	−6.17
	ECM		
lags	2	2	2
EC_{t-1}	−0.237 (0.069)	−0.326 (0.083)	−0.622 (0.100)
R^2	0.56	0.58	0.66

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 unc_{1t} 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

[38] The factor unc_{2t} 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.

[39] The 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).[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
first observation 1975Q1

end of sample	2004Q4	2001Q3	1998Q4	2004Q4	2001Q3	1998Q4
			long-run relation			
y_t	1.226 (0.200)	1.135 (0.122)	1.023 (0.030)	1.144 (0.053)	1.114 (0.048)	1.031 (0.024)
oc_t	-4.587 (2.953)	-4.232 (1.714)	-4.465 (0.371)	-3.024 (0.664)	-3.453 (0.595)	-4.454 (0.303)
unc_{1t}	-	-	-	1.552 (0.224)	1.037 (0.278)	-0.070 (0.205)
unit root	-1.68	-1.78	-4.31	-4.06	-3.80	-5.62
			ECM			
lags	2	2	2	2	2	2
EC_{t-1}	-0.047 (0.023)	-0.072 (0.039)	-0.326 (0.069)	-0.130 (0.051)	-0.183 (0.067)	-0.573 (0.094)
R^2	0.51	0.51	0.59	0.54	0.55	0.66

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

[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.

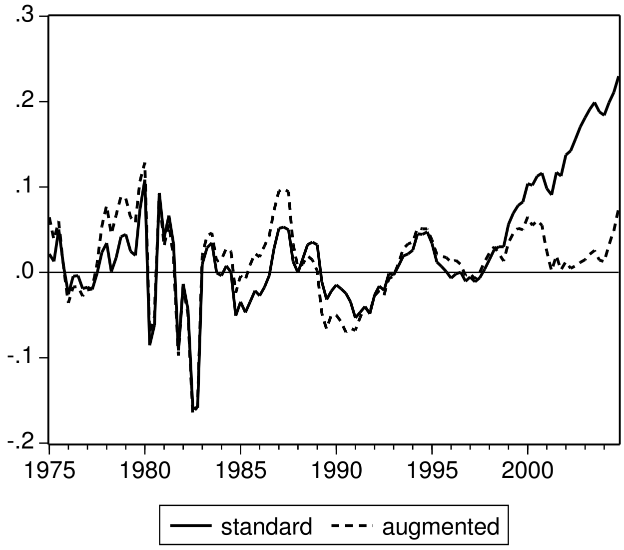

Figure 10: Monetary overhang, M2M

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.

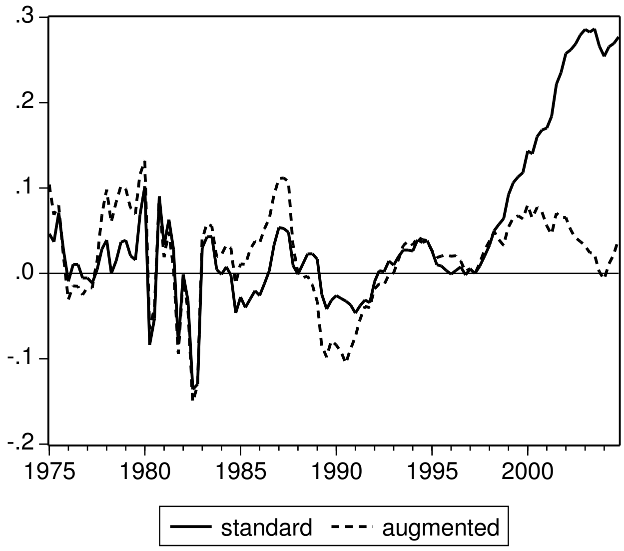

Figure 11: Monetary overhang, MZM

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.

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