**Kiel Institute for World Economics**
Duesternbrooker Weg 120
24105 Kiel (Germany)


**Kiel Working Paper No. 1179**

# Is European Money Demand Still Stable?

by

**Kai Carstensen**

Revised version, March 2004

# Is European money demand still stable?

*Kai Carstensen*[*]

*Kiel Institute for World Economics, Duesternbrooker Weg 120, 24105 Kiel, Germany*

*E-mail:* `carstensen@ifw.uni-kiel.de`

**Abstract**

This paper analyzes the question whether money demand in the Euro area has undergone a structural change in recent time when M3 money growth has considerably overshot the reference value set by the European Central Bank (ECB). It is found that conventional specifications of money demand have in fact become unstable while specifications which are augmented with equity returns and volatility remain stable. Using such an augmented specification, it turns out that the excessive M3 growth rates can largely be attributed to the stock market downswing and do not put a measurable threat to price stability.

JEL code: E41

Keywords: Money demand, EMU, excess liquidity.

## 1 Introduction

On 8 May 2003 the ECB announced a revision of its monetary policy strategy (ECB, 2003b). By most observers, the revision was interpreted as a weakening of the first pillar (de Grauwe, 2003, Belke et al., 2003). It might in part be motivated by the fact that M3 reference growth rates have been continuing to exceed the reference value of 4.5 percent by more than 2.5 percentage points since the end of 2001. At the same time, the ECB lowered its key interest rate from 4.25 percent on 31 August 2001 to 2 percent on 6 June 2003 although the monetary developments suggested opposite action.

The ECB explains the strong money growth with portfolio shifts from equities

[*]I wish to thank, without implicating, Donald Andrews, Joachim Scheide, Beatrice Weder and the participants of the *Freitagsseminar* at the Bundesbank for helpful comments.

to safe and liquid assets which are induced by the recent stock market downswing and the increased financial uncertainty and will be reversed once stock prices rise again and uncertainty diminishes (see, e.g. ECB, 2003a). From this perspective, the recent money growth does not seem to pose a particular threat to price stability. It might, however, indicate that the relationship between money and prices has become unstable and, hence, money growth is not a well-suited tool to analyze prospective inflation and support monetary policy decisions. It would then be only natural that the ECB reduced the weight of the second pillar.

Because it is generally assumed that money and prices are related via a money demand function, the preceding discussion raises the question whether European money demand has recently become unstable. There is a large number of papers which deal with estimating money demand functions of the European Monetary Union (EMU) and testing their stability. Most of them exclusively use synthetic data for the pre–EMU period[1] or up to the first year of EMU[2] and cannot reject stability. Extending the data set until the third quarter of 2001, Kontolemis (2002) finds evidence for an instability of the conventional money demand function at his very last observation due to the strong growth of M3 beginning in this period.

In a comprehensive stability analysis Bruggeman et al. (2003) apply the fluctuation and Nyblom–type stability tests proposed by Hansen and Johansen (1999) and obtain mixed results but finally conclude that there are some specifications of money demand which seem stable. However, since their data set ends with the the fourth quarter of 2001 and the excessive money growth did not start before the second quarter 2001, it is well possible that their limited data set prevented the statistical tests from indicating non–stability.

This paper adds to the literature by, first, using an updated data set from the

[1] E.g., Gottschalk, 1999, Hayo, 1999, Bruggeman, 2000, Clausen and Kim, 2000, Coenen and Vega, 2001, Funke, 2001, Müller and Hahn, 2001, and Golinelli and Pastorello, 2002.

[2] Brand and Cassola, 2000, and Calza et al., 2001.

first quarter of 1980 until the second quarter of 2003. Consequently, there are more observations with excessive money growth at the end of the sample available. Second, we supplement existing stability tests with a new family of stability tests proposed by Andrews and Kim (2003) which perfectly fits our purpose because it is designed to detect a breakdown of cointegration at the end of a sample. In addition, we also test for short–run instability using a similar test put forward by Andrews (2003) for stationary environments. Since we find conventional money demand specifications to become unstable in 2001, we specify money demand functions augmented with stock market variables which exhibit structural stability and can be used to quantitatively assess the importance of stock market developments on M3 growth rates. It turns out that the recent excess M3 growth rates do not pose a threat to price stability.

The remainder of this paper is organized as follows. In Section 2 we briefly review the literature on EMU money demand function. In Section 3 we outline the end–of–sample stability tests proposed by Andrews (2003) and Andrews and Kim (2003). The empirical test results and alternative specifications are described in Section 4 while the policy implications of the findings are discussed in Section 5. Finally, Section 6 concludes.

## 2 Specifications of the European money demand function

This section gives a brief overview over different specifications used in the literature for the European money demand function. An extensive review is provided by Golinelli and Pastorello (2002). In this paper, we concentrate on demand for real M3 $mp_t = m_t - p_t$ which is usually assumed to depend on real GDP $y_t$, its own rate $r_t^o$, a short term interest rate $r_t^s$, a long term interest rate $r_t^l$ and the inflation rate $\Delta p_t$,

$$mp_t = \beta_1 y_t + \beta_2 r_t^l + \beta_3 r_t^s + \beta_4 r_t^o + \beta_5 \Delta p_t + u_t. \qquad (1)$$

This full specification is generally not estimable due to collinearity between the regressors, notably the interest rates. Typical specifications used in the recent literature on European money demand are presented in Table 1. They differ in the use of the interest rates and the inflation rate. In specification S1 money demand depends on GDP and the long–term interest rate as a measure of the opportunity costs of holding money. This specification is estimated by, e.g., Golinelli and Pastorello (2002). Since M3 includes a number of interest–bearing securities, it is often argued that one should also consider a measure of the own rate of M3. Therefore, some authors include both the long–term rate as a measure of the opportunity costs and the short–term rate as a proxy for the own rate (Gottschalk, 1999, Clausen and Kim, 2000, Müller and Hahn, 2001). Following Calza et al. (2001), we construct a direct measure of the own rate and include the spread between the long–term rate and the own rate together with GDP in our specification S2. Augmenting this specification with the inflation rate yields the model estimated by Coenen and Vega (2001) which makes up our specification S3. Instead of including the long–term interest rate as a measure of the opportunity costs of M3, some authors propose including the short–term interest rate. This gives us the specifications S4 to S6: In specification S4, we include GDP and the short–term interest rate, in specification S5 we include GDP and the spread between the short–term interest rate and the own rate, and in specification S6 we additionally include the inflation rate. Furthermore, we use a variant where money demand solely depends on GDP and the inflation rate (specification S7). This specification has been successfully applied to German money demand by Wolters et al. (1998) and Lütkepohl and Wolters (2003) and might be a viable alternative for the EMU area of which Germany is the largest member country. Finally, one might assume that the negative trend in M3 income velocity which is analyzed by Brand et al. (2002) can solely be explained by an income elasticity of money demand greater than 1. In this case, money demand depends on GDP only (specification S8).

Most of these specifications are found by other authors to be stable and cointegrated in earlier sample periods which generally end before or with the beginning of EMU. Our aim is to test the hypothesis of structural stability for each of the specifications in an updated sample comprising observations from 1980Q1 until 2003Q2.[3] To this end, we use the data set published by Calza et al. (2001) which contains data for M3, GDP, the GDP deflator, the long–term and the short–term interest rate, and the own rate of M3 from 1980Q1 until 1999Q4, but extend it until 2003Q2 as explained in the appendix. All variables except for the interest rates are given in logs.

## 3 Tests for end–of–sample stability

It has long been recognized that the variables entering the money demand function (1) can best be modelled as integrated I(1) processes. Therefore, stability of the money demand function requires as a minimum that (1) constitutes a cointegration relationship. It is by now a well–established empirical finding that the European money demand function in fact constitutes a cointegration relationship at least for the sample from 1980Q1 until 1998Q4 which we will call the baseline sample. It is for this reason that we do not replicate a comprehensive cointegration analysis for this sample but refer to the work by, inter alia, Calza et al. (2001), Brand and Cassola (2000) and Bruggeman et al. (2003). Instead, we condition on the assumption that the money demand function in its various specifications represents a stable cointegration relationship in the baseline sample and test whether this stability has been lost since then, especially since M3 started its excessive growth.

If the long–run or cointegration parameters are constant, the model exhibits long–run stability. If, in addition, also the short–run parameters, i.e., the parameters for lagged differences of the variables which are used to model transitory fluctuations, are found to be constant, the model can be said to exhibit full structural stability.

[3] In the estimation exercises below, the first four observations are reserved as pre-sample values.

Due the superconsistency of estimators for cointegration parameters, we can split the problem of stability testing into two sub–problems. In a first step, we analyze the stability of the cointegration parameters. If stability is found or restored by, e.g., inclusion of dummy variables, we can analyze the stability of the short–run parameters in a second step taking the superconsistently estimated cointegration parameters as given. Therefore, we need two end–of–sample stability tests, one for cointegrating regressions and one for stationary regressions. We use the stability tests put forward by Andrews and Kim (2003) for cointegrating regressions and Andrews (2003) for stationary regressions. These tests are generalizations of the well-known Chow stability test and are easy to compute. Moreover, critical values and $p$–values can be obtained from a parametric subsampling which circumvents the use of asymptotic distributions. This is particularly important if the typical assumption needed to derive the asymptotic distribution for a structural–break model, namely that the lengths of the pre–break and post–break periods are of a fixed proportion even asymptotically, is deemed unrealistic. Instead, it is assumed that the post–break sample is of fixed and finite length.

## 3.1 Stability tests in cointegrating regressions

An end–of–sample stability test for cointegrating regressions is proposed by Andrews and Kim (2003) who call it a cointegration breakdown test. Splitting the sample of size $t = 1, \ldots, T + m$ into the first $T$ and the last $m$ observations, they start from the linear model

$$y_t = \begin{cases} x_t'\beta_0 + u_t, & t = 1, \ldots, T \\ x_t'\beta_t + u_t, & t = T+1, \ldots, T+m, \end{cases} \tag{2}$$

where the regressors are allowed to be linear combinations of integrated I(1) random variables, stationary random variables and deterministic variables. They test the null hypothesis that the model is stable and cointegrated, i.e., $\beta_0 = \beta_t$ for all $t =$

$T+1,\ldots,T+m$ and $u_t$ is stationary for all $t=1,\ldots,T+m$, against the alternative hypothesis that either $\beta_0 \neq \beta_t$ for some $t \in \{T+1,\ldots,T+m\}$ or the distribution of $\{u_{T+1},\ldots,u_{T+m}\}$ differs from the distribution of $\{u_1,\ldots,u_m\}$. In particular, a shift in the parameter vector $\beta_0$ to $\beta_t$ or a shift in the distribution of $u_t$ from being stationary to being integrated I(1) should cause the null hypothesis to be rejected. Both cases can be interpreted as a cointegration breakdown. Note that the setup does not require the break occur exactly at time $T+1$ but rather in the interval $\{T+1,\ldots,T+m\}$.

The first family of tests is of a Chow–type. Applying, e.g., ordinary least squares (OLS), fully modified least squares (FM–OLS) proposed by Phillips and Hansen (1990) or full information maximum likelihood estimation (FIML) proposed by Johansen (1988, 1991) to model (2) for the first subsample $t=1,\ldots,T$, gives rise to the estimator $\hat{\beta}_{1—T}$. In the next step, this estimator is used to compute the prediction errors

$$\hat{u}_t = y_t - x_t'\hat{\beta}_{1—T}, \qquad t = T+1,\ldots,T+m, \tag{3}$$

from which the sum–of–squares statistic

$$P_a = \sum_{t=T+1}^{T+m} \hat{u}_t^2 \tag{4}$$

is calculated. This test statistic is supplemented by two similar ones, $P_b$ and $P_c$, which are based on the estimators $\hat{\beta}_{1—(T+[m/2])}$ and $\hat{\beta}_{1—(T+m)}$, respectively, but are otherwise equal.[4]

To determine critical values and $p$–values, Andrews and Kim propose the use of a parametric subsampling technique instead of large–sample asymptotics. Under the null hypothesis, the stationarity assumption for $u_t$ ensures that the distribution of

[4] Note that $[m/2]$ denotes the smallest integer greater than or equal to $m/2$.

the statistic

$$P_1(\beta_0) = \sum_{t=1}^{m} (y_t - x_t'\beta_0)^2 \tag{5}$$

converges to the distribution of $P_a$ because $\hat{\beta}_{1-T}$ used to compute $P_a$ converges in probability to the true parameter vector $\beta_0$. Since the random variables

$$P_j(\beta_0) = \sum_{t=j}^{j+m-1} (y_t - x_t'\beta_0)^2, \qquad j = 1, \ldots, T-m+1, \tag{6}$$

are stationary and ergodic, the empirical distribution function of $P_j(\beta_0)$, $j = 1, \ldots, T-m+1$, is a consistent estimator for the distribution function of $P_1(\beta_0)$ and, hence, $P_a$. However, $\beta_0$ in $P_j(\beta_0)$ is unknown, so it must be estimated. To mimic the property of the $P_a$ statistic that the estimation sample $t = 1, \ldots, T$ and the prediction sample $t = T+1, \ldots, T+m$ are non–overlapping, Andrews and Kim suggest to evaluate $P_j$ at the "leave–$m$–out" estimator $\hat{\beta}_{(j)}$ which uses the observations $t = 1, \ldots, T$ with $t \neq j, \ldots, j+m-1$. Given the empirical distribution function of $P_j$, the computation of critical values and $p$–values is straightforward. For $P_b$ the same distribution function is applied, while for $P_c$ the "leave–$m$–out" estimator used to compute $P_j$ is replaced by a "leave–$[m/2]$–out" estimator.

The second family of tests is motivated by a locally best invariant (LBI) test for the presence of unit root disturbances in the second subsample $t = T+1, \ldots, T+m$. A test statistic analogous to $P_a$ is defined by the weighted sum

$$R_a = \sum_{i=T+1}^{T+m} \sum_{j=T+1}^{T+m} \min\{i-T, j-T\}\, \hat{u}_i \hat{u}_j = \sum_{i=T+1}^{T+m} \left( \sum_{j=i}^{T+m} \hat{u}_j \right)^2, \tag{7}$$

where $\hat{u}_i$, $i = T+1, \ldots, T+m$, denote the prediction errors given in (3). Again, two additional test statistics $R_b$ and $R_c$ are proposed. They are computed in the same fashion as $P_b$ and $P_c$, i.e., using the estimators $\hat{\beta}_{1-(T+[m/2])}$ and $\hat{\beta}_{1-(T+m)}$, respectively, but apply a weighted sum like (7) to the resulting prediction errors instead of a sum of squares like (4). Critical values and $p$–values are also calculated analogously.

Andrews and Kim report an extensive simulation study from which they conclude that the $P_a$ and $R_a$, and, to a lesser extent, the $P_b$ and $R_b$ tests over–reject the true null hypothesis of structural stability. Therefore, especially the former two tests are not recommended. On the other hand, the $R_c$ test slightly under–rejects the true null. However, particularly $R_c$ but also the $P_c$ tests are found to possess the best power properties both against the alternative of a shift in the parameter vector and a change of the error distribution from being stationary to being integrated I(1). For this reason we use these two tests in the empirical analysis of Euro area money demand. Note that the $P_c$ test is designed for the alternative hypothesis of parameter instability whereas the $R_c$ is designed for the alternative hypothesis of the disturbances changing from being stationary to being integrated I(1). Surprisingly, the former test seems to possess even more power against the latter hypothesis and vice versa. Thus, we will use both the $P_c$ and the $R_c$ statistics to test for stationary disturbances and parameter stability.

## 3.2 Stability tests in stationary regressions

The end–of–sample stability tests for stationary regressions proposed by Andrews (2003) are a direct generalization of an F test for structural change and similar to the cointegration breakdown tests described above. In contrast to the F test, both lagged endogenous explanatory variables and non–normal, heteroskedastic and auto-correlated disturbances are allowed. The model setup (2) is now used to test the null hypothesis that the model is stable, i.e., $\beta_0 = \beta_t$ for all $t = T+1, \ldots, T+m$ and the distribution of all $u_i$, $i = T+1, \ldots, T+m$, equals the distribution of $u_i$, $i = 1, \ldots, T$, against the alternative hypothesis that either $\beta_0 \neq \beta_t$ for some $t \in \{T+1, \ldots, T+m\}$ or the distribution of some $u_i$, $i = T+1, \ldots, T+m$, differs from the distribution of

$u_i$, $i = 1, \ldots, T$.

In a way similar to the cointegration breakdown tests, Andrews defines several slightly different stability tests for stationary regressions but concludes from a simulation study that one specific tests unanimously dominates all its competitors. Only this test will be sketched in the following and, subsequently, used to determine the short–run stability of the EMU money demand function.

In a first step, a GLS transformation is applied to the model in order to restore uncorrelated and homoskedastic disturbances. To this end, the error covariance matrix is estimated as

$$\hat{\Sigma} = (T+1)^{-1} \sum_{t=1}^{T+1} \hat{U}_t \hat{U}_t' \tag{8}$$

where $\hat{U}_t = (\hat{u}_t, \ldots, \hat{u}_{t+m-1})'$ and $\hat{u}_t = y_t - x_t' \hat{\beta}_{1—(T+m)}$. Pre–multiplying the model in the post–break sample by $\hat{\Sigma}^{-1/2}$ and defining $V_t = \hat{\Sigma}^{-1/2} U_t$, $\bar{Y}_t = \hat{\Sigma}^{-1/2}(y_t, \ldots, y_{t+m-1})'$ and $\bar{X}_t = \hat{\Sigma}^{-1/2}(x_t, \ldots, x_{t+m-1})'$ then yields a model with i.i.d. disturbances

$$\bar{Y}_t = \bar{X}_t \beta + V_t, \tag{9}$$

on which the stability test is based. Given the length of the post–break sample $m$ is larger than the number of regressors so that $\bar{X}_{T+1}' \bar{X}_{T+1}$ is invertible, the test statistic is finally given by

$$S_d = \hat{V}_{T+1}' \bar{X}_{T+1}' \left( \bar{X}_{T+1}' \bar{X}_{T+1} \right)^{-1} \bar{X}_{T+1} \hat{V}_{T+1}, \tag{10}$$

where $\hat{V}_{T+1} = Y_{T+1} - X_{T+1} \hat{\beta}_{1—T+m}$. Otherwise it is simply

$$S_d = \hat{V}_{T+1}' \hat{V}_{T+1}. \tag{11}$$

Critical values and $p$–values are estimated as described for the cointegration breakdown $P_c$ test, i.e., using the "leave–[m/2]–out" estimator.

### 3.3 Conventional tests for structural breaks

We also consider several structural break tests used elsewhere in the literature on EMU money demand. To analyze long–run stability, we apply Nyblom–type tests designed for cointegrated VAR models (SupQ and MeanQ statistics, Hansen and Johansen, 1999) and for cointegrating regressions ($L_c$ statistic, Hansen, 1992b). To analyze short–run stability, we report the fluctuation test of Ploberger et al. (1989) and the Nyblom–type test of Hansen (1992a). Usually, it is deemed a particular strength of these tests that they do not require a prespecified break date but test the null hypothesis of structural stability against the alternative that there is a structural break at some unknown point in the sample. For the present situation, this strength may turn out as a weakness because there are two known dates in the sample where instability may show up, namely the start of EMU in 1999Q1 and the start of excessive M3 growth around 2001Q4. While the former date is clearly identified exogenously, one can argue that the latter is not; neither can it be unambiguously fixed to exactly one quarter. On the other hand, there is ample evidence from newspapers, ECB reports, commentators and other sources that something in fact was going on around this date. Therefore, tests which use 2001Q4 as a known break date are probably better suited than tests which abstract from any a–priori knowledge and, thus, lack power.

## 4 Empirical results

In this section we present and discuss the results of the stability tests applied to the eight specifications of the Euro area money demand function. We proceed as follows. First, the appropriateness of the specifications for the baseline sample 1980Q1 to 1998Q4 is examined. To this end, tests for cointegration and structural stability are carried out and the parameters are estimated by means of OLS, FM–OLS and FIML. Subsequently, recursive parameter estimates and the cointegration–breakdown tests of Andrews and Kim (2003) are reported for the quarters 1999Q1 to 2003Q2 in order to test for long–run stability in this period. Since long–run stability is rejected, specifications augmented with stock market variables are proposed in order to restore stability. For these specifications, the cointegration and stability analysis is replicated. It turns out that it is in fact possible to specify stable and plausible long-run relationships. Finally, following the suggestion by Hansen (1992a), for each specification an error–correction models is put up, taking the cointegration parameters estimated by FM–OLS as given, and used to test for short–run stability.

## 4.1 Cointegration and stability in the baseline sample

In a first step, the cointegration properties in the baseline sample are analyzed by means of Bartlett corrected trace tests (Johansen, 2002) which are reported in Table 2 together with $p$–values derived from the asymptotic distribution.[5] At the ten percent level, there is one cointegration relationship between the variables of specifications S1, S3, S4, S5, S6 and S7 whereas there is no cointegration between the variables of specifications S2 and S8. While the latter result may be due to the power problems of the trace test, it casts some doubts on the appropriateness of specifications S2 and S8. On the other hand, Müller and Hahn (2001), *inter alia*, find specification S2 to be cointegrated. Thus for the time being, it is nevertheless assumed that each specification represents a cointegration relationship.

The estimated parameters of the eight specifications S1 to S8 of the money demand function (1) are displayed in Table 3. The income elasticity $\beta_1$ is estimated highly significant and remarkably stable as slightly below 1.4 over most specifications and estimation methods. This stability is also reported by Brand et al. (2002) who use

[5] Using bootstrapped $p$-values instead does not lead to any important changes.

data until 2001Q2. Significant and plausible values for the remaining parameters are only found for specifications S4, S5, S7 and S8. In specification S1, the semi–elasticity of the long–term rate has the expected negative sign but is not significant when using FIML. In specification S2, the coefficient of the spread between the long–term rate and the own rate is either insignificant (FM–OLS) or implausible (FIML). This does not change when adding the inflation rate in specification S3.[6] In specification S4, the semi–elasticity for the short–term rate has the expected negative sign and is significant. This holds even better if the spread between the short–term rate and the own rate is used instead in specification S5. However, adding the inflation rate in specification S6 again asks too much of the data leading to insignificant or implausible estimates. Dropping all interest rates in specification S7 leads to significant and plausible estimates of the long-run parameters for GDP and inflation. The same holds for specification S8 where money demand solely depends on GDP.

In a next step, it is analyzed whether each of these relationships is stable over the baseline sample. To this end, Nyblom–type stability tests are used because we have no prior for any particular break date. In Table 4, both the $L_c$ test of Hansen (1992b) and the SupQ and MeanQ tests of Hansen and Johansen (1999) are reported. At the 10 percent level, the $L_c$ test rejects stability for specification S1 while the SupQ and MeanQ tests reject stability for specification S2. All other specifications turn out to be stable in the baseline sample. This result is expected because the stability of Euro area money demand in early samples is well documented in the literature (see also the literature review by Calza and Sousa, 2003).

Taking the outcomes of the cointegration and stability tests, and the parameter estimates together, specifications S4, S5 and S7 particularly qualify for further

[6] For specification S3 the results stand in contrast to the results presented by Coenen and Vega (2001) who use the long–run solution of an autoregressive distributed lag (ADL) model to estimate the cointegration parameters. However, taking their data set and employing OLS and FM–OLS we obtain insignificant estimates for the parameter of the interest rate spread. Applying the ADL method to our data set also yields an insignificant weight of the interest rate spread. From this we conclude that their specific data set and method play an important role to obtain plausible results.

consideration because they exhibit stable and plausible parameter estimates and a cointegration rank of 1. Of these three candidates, we prefer S5 because the opportunity costs of holding money are modelled most appropriately in this specification as argued by Calza et al. (2001). Nevertheless, we will continue to report test results for all specifications to circumvent any selectivity bias in our conclusions.

## 4.2 Tests for end–of–sample stability of the long–run relationships

To obtain a first impression of the stability properties of long–run EMU money demand, we present recursive FM–OLS parameter estimates together with 90 percent confidence intervals for the specifications S4, S5 and S7, see Figure 2. There is little variation of the estimated parameters during the time from 1995Q1 to 2001Q4. Only at the end of the sample there is a strong sign of instability. However, the lack of statistical guidance as to whether these parameter shifts are really significant, is the main drawback of recursive parameter estimates.

Therefore, in a second step, we test for cointegration breakdown as outlined above using OLS, FM–OLS and FIML to estimate the long–run relationships. We consider as potential breakpoints the begin of EMU, 1999Q1, and the begin of excessive money growth, 2001Q4. The first breakpoint is analyzed in the sample 1981Q1 to 2001Q3, i.e., before excessive money growth started, and in the full sample 1981Q1 to 2003Q2. The second breakpoint is analyzed in the full sample only. The simulated $p$–values of the $P_c$ and the $R_c$ tests are displayed in Table 5.

As a general result except for specification S6 estimated by FIML, all test statistics and specifications indicate that there is no sign of instability at the begin of EMU if only the sample until 2001Q3 is considered. However, if the sample is extended until 2003Q2, stability is generally rejected. This suggests that it is not the begin of EMU which causes instability, but some change after 2001Q3. In fact, stability since 2001Q4 in the full sample is clearly rejected. From this we conclude that a cointegration breakdown occurred in the period starting 2001Q4 and, thus, coincides with the period of excessive money growth.[7]

### 4.3 Modelling the long–run structural change

The test results presented in the previous section indicate that long–run structural stability of the EMU money demand function probably failed since the end of 2001. It is therefore of interest to explain, or even model, this structural change. The coincidence of excessive money growth and stock market turmoil suggests that there might be a relationship between money demand and stock prices. Friedman (1988) argues that there are several ways by which stock markets should affect money demand: Real stock prices should have a positive wealth effect, stock returns should have a negative substitution effect and stock market risk should have a positive risk-avoidance effect given risk-averse agents. For the US, he can empirically confirm these claims. Choudhry (1996) and Carpenter and Lange (2002) also find evidence in favor of a long–run influence of stock market variables on US money demand. Caruso (2001) extends this work to a panel of 25 countries and obtains a significant wealth effect. For EMU money demand, Kontolemis (2002) using data up to 2001Q3 finds a significant long–run influence of stock prices while Bruggeman et al. (2003) using data up to 2001Q4 do not obtain significant parameters for either real stock prices nor for stock market volatility. The latter result stands in contrast to the argument of the ECB (2003a) that the increased uncertainty in equity markets has led to portfolio shifts from equities to safe and liquid assets, which are part of M3, and hence to the excessive M3 growth. Moreover, Cassola and Morana (2002), using data up to

[7] The main exceptions are specifications S2 and S6 for which stability cannot be rejected in the full sample if the cointegration breakdown tests are based on FIML estimation. For specification S2 this can be explained by the finding that stability in the baseline sample is rejected while the cointegration breakdown tests condition on stability before the supposed break date. Similarly, stability of specification S6 after the start of EMU is rejected in the sample until 2001Q3. Moreover, the parameters in both specifications take very implausible values which makes it difficult to interpret them as money demand relationships.

2000Q4, find that real stock prices play an important role in the monetary transmission mechanism. It is therefore an unresolved issue whether stock market variables can account for the instability in the EMU long–run money demand function. For our extended data set, we analyze this question using three different stock market variables, namely, real stock prices $sp_t$, stock returns $r_t^e$ and stock market volatility $v_t$. The construction of these variables is explained in the appendix.

In a first step, we use three unit root tests to determine whether the stock market variables are integrated which is a necessary condition for them to enter the cointegration relationship. The results of the augmented Dickey–Fuller (ADF) test, the Phillips–Perron (PP) tests and the DFGLSu test (Elliott, 1999) are displayed in Table 6. We cannot reject the null hypothesis of nonstationarity in any case.[8] Therefore, we proceed under the assumption that the stock market variables are nonstationary.

We now augment the conventional money demand specifications with some of the stock market variables. Three different augmentations are found to yield sensible results. First, the specifications S1a to S8a are augmented with the spread between equity returns and the own rate $r_t^e - r_t^o$ as a measure of stock market opportunity costs.[9] Second, the specifications S1b to S8b additionally include the real stock price index $sp_t$. Finally, the specifications S1c to S8c include the spread between equity returns and the own rate $r_t^e - r_t^o$ and stock market volatility $v_t$. Specifications which include all three variables do not yield significant parameter estimates and are, thus, not reported.

The FM–OLS and FIML estimation results are presented in Tables 7 and 8,

[8] In a very strict sense, it may be difficult to defend the nonstationarity result for volatility if it is assumed that volatility cannot become arbitrarily large. However, this problem equally applies to many other variables, notably interest rates and inflation rates, for which the nonstationarity assumption is generally maintained. In our view it is sufficient that the series at hand behaves like a nonstationary variable in the given sample. For our volatility measure, this is certainly the case, see Fig. 1. Moreover, volatility is constructed from a GARCH model with nearly nonstationary conditional variance equation.

[9] To mimic the conventional specifications S1 and S4 which use a single long-term and short-term interest rate, respectively, instead of the spread to the own rate, we replace $r_t^e - r_t^o$ with equity returns $r_t^e$ in specifications S1a and S4a.

respectively. As expected, the spread $r_t^e - r_t^o$ has a negative impact on long–run money demand indicating a substitution effect, whereas real stock prices $sp_t$ and volatility $v_t$ positively affect long–run money demand indicating a wealth and risk-aversion effect, respectively. It strikes the eye that the preferred specification S5 exhibits significant and plausible parameter estimates in all extensions and for both estimations techniques. The same holds, with few exceptions, for S4 and S7 which qualified as good candidates in the baseline sample, while the other specifications in most cases fail to produce satisfying results.

Using the money demand functions augmented with stock market variables is only a viable alternative to conventional specifications if they solve the instability problem. To this end, we again carry out a number of tests. First, Bartlett corrected trace tests reported in Table 9 indicate that it is generally difficult to find cointegration for specifications S1a to S8a and S1b to S8b. While this may be due to the well-known power problems of the trace test, it reduces the attractiveness of these specifications. In contrast, specifications S1c and S3c to S7c clearly have cointegration rank 1. Next, we analyze long–run stability given a cointegration rank of 1. In Table 10, the simulated $p$–values of the cointegration–breakdown tests are presented. At the 5% level, stability can be accepted with the only exceptions being specifications S3a, S3b, S7a and S7b for the breakpoint 2001Q4. In addition, Nyblom–type tests are reported in Table 11. Except for specification S1c for which the $L_c$ and SupQ statistics are significant at the 10 percent level, all specifications exhibit long–run stability.

From these results we conclude that augmenting conventional money demand functions with stock market variables can help restore stable long-run relationships, especially when the spread between equity returns and the own rate, $r_t^e - r_t^o$, and stock market volatility, $v_t$, are added to the conventional variables. Among the variety of different possible specifications, S5c is particularly appealing both theoretically, because the opportunity costs are measured correctly, and empirically, because this specification passes all relevant tests without problem. This is reassured by recursive parameter estimates of this specification (Figure 3) which do not show the break that was evident for specification S5 (Figure 2). Therefore, we will use S5c for policy analysis in section 5.

### 4.4 Tests for short–run stability

Before turning to policy implications of the augmented money demand functions, we analyze their short–run stability properties. For brevity, we concentrate on specifications S1c to S8c which performed best with the cointegration tests. In a first step, Ploberger et al. (1989) fluctuation tests are applied to each equation of the VAR models used to perform FIML estimation of the cointegration parameters. The results displayed in Table 12 do not indicate instability at the 5 percent level.

In addition, for each specification we set up parsimonious single–equation error–correction models (ECM) from which insignificant variables are excluded. In order to enhance the model fit, a large number of possibly lagged explanatory variables are allowed to enter the model: the cointegration residuals $\hat{u}_t^s$ ($s = \text{S1c}, \ldots, \text{S8c}$), $\Delta mp_{t-1}$, $\Delta y_t$, $\Delta r_t^l$, $\Delta r_t^s$, $\Delta r_t^o$, $\Delta^2 p_t$, $\Delta sp_t$, $\Delta r_t^e$, $\Delta v_t$ and $\Delta oil_t$ which is an oil price index described in the appendix. Note that the cointegration residuals $\hat{u}_t^s$ are estimated by FM–OLS which is asymptotically justified by the superconsistency of the estimated cointegration parameters. It turns out that the following lag structure with 11 free parameters yields both uncorrelated disturbances and highly significant parameter estimates for all specifications:

$$\begin{aligned}\Delta mp_t &= \gamma_0 + \gamma_1 \hat{u}_{t-1}^s + \gamma_2 \Delta mp_{t-1} + \gamma_3 \Delta r_{t-1}^l + \gamma_4 \Delta r_{t-4}^s + \gamma_5 \Delta r_{t-2}^o \\ &+ \gamma_6 \Delta^2 r_{t-3}^o + \gamma_7 \Delta^2 p_{t-2} + \gamma_8 \Delta^2 v_t + \gamma_9 \Delta sp_{t-1} + \Delta oil_{t-1} + \varepsilon_t. \qquad (12)\end{aligned}$$

Note that the inclusion of the contemporaneous value of $v_t$ is valid because tests indicate that $v_t$ is weakly exogenous in each specification.

The $p$–values of the end–of–sample stability test $S_d$ for stationary regressions are presented in Table 12. Over all specifications, the estimated $p$–values are far above the 5% level both for the breakpoints 1999Q1 and 2001Q4. Consequently, the null hypothesis of structural stability cannot be rejected. As a supplement, we compute Nyblom–type tests for structural stability in stationary regressions (Hansen, 1992a). Again, stability cannot be rejected. We conclude that it is possible to specify augmented money demand functions for the Euro area which exhibit both long–run and short–run stability and, thus, are well suited for policy analysis.

## 5 Policy Implications

In several Monthly Bulletins (e.g. ECB, 2003a) the ECB argues that portfolio shifts from stock markets to safe instruments which are part of M3 have caused the excessive growth rates of nominal M3. Using the estimated money demand function (specification S5c), it is possible to evaluate the importance of the stock market developments. To this end, we decompose the annual long–run real money demand changes into the contributions caused by changes in the individual explanatory variables (Table 13). Until 2000, the conventional variables, $y_t$, $r_t^s$ and $r_t^o$, account for the major part of the long–run money demand changes. However, since 2001, the stock market variables play the most important role. Consider, e.g., the 2002 increase in real money demand of 5.3 percent. Only 1.6 percent are caused by the conventional variables while the stock market variables account for 3.6 percent. This underlines the economic importance of the recent stock market developments for money demand in the Euro area.

We can go a step further and analyze the *ceteris paribus* cumulated effect of the stock market variables on money demand since the all-time high in 2000Q3. Thereby, we can quantify the share of money growth which is due to the stock market downturn. To this end, we use the estimated single-equation error–correction model for specification S5c. To calculate the stock market effect we dynamically simulate the model conditional on, first, the observed paths of stock market yields and volatility (reference simulation) and, second, the assumption that these variables have stayed constant since 2000Q3 (“constant stock market” simulation). For all other variables, the observed paths are used. The difference between the two simulations can then be exclusively traced back to the stock market influence on money demand.[10]

The simulated annual M3 growth rates $m_t^{sim} - m_{t-4}^{sim} = mp_t^{sim} - mp_{t-4}^{sim} + p_t - p_{t-4}$ are displayed in the upper panel of Figure 4 together mit actual M3 growth rates. Actual and simulated M3 growth rates are, by assumption, identical up to 2000Q3. Afterwards, M3 growth rates of the reference simulation track actual M3 growth rates reasonably well whereas M3 growth rates of the “constant stock market” simulation lie considerably below, particularly at the end of the sample where the reference simulation and the “constant stock market” simulation yield growth rates of about 8 percent and 4.5 percent, respectively. The difference of 3.5 percent growth rate can be attributed to the stock market downswing.

The levels effect of the stock market downswing is displayed in the lower panel of Figure 4. In 2003Q2, the downswing caused an increase in M3, measured as the difference between the reference simulation and the “constant stock market” simulation, of roughly 310 Billion Euro. As a comparison, the ECB (2003a) calculates an effect of 180 to 250 Billion Euro for the shorter period 2001Q2 to 2003Q1 which is of a comparable magnitude. Note that for the period 2001Q2 to 2003Q1, we obtain an effect of 240 Billion Euro which is at the upper limit of the interval estimated by the

[10] Note that this exercise does not provide any information about the sources of the changes of the stock market variables. Moreover, we do not change the paths of the remaining explanatory variables. The experiments therefore neglect any interrelations between these variables and the stock market developments. In particular, we do not disentangle and identify structural shocks and their impacts on money demand (for a recent paper, see Cassola and Morana, 2002). As a consequence, we do not analyze the effects of, say, exogenous stock market shocks on the whole economy and, as a part of it, on money demand. Nevertheless, the experiments are still informative because they help answer the question how the stock market developments *ceteris paribus* have affected money demand.

ECB. The results suggest that a considerable, and in 2003Q2 still rising, portion of money growth is due to the stock market downswing and increased volatility.

For monetary policy, the interesting question is whether the unusually high money growth rates induce any excess liquidity with potentially inflationary consequences. The answer to this question depends on how excess liquidity should be measured. If one takes the reference value seriously, it implies that the deviations of money growth rates from 4.5% accumulate over time because M3 departs more and more from the reference trend line with slope 0.045. Eventually, this should lead to rising inflation. However, stock market considerations did not play any role in the derivation of the money growth reference value. Therefore, this measure might be misleading in the current situation. A more appropriate measure would be the money growth rate adjusted for the stock market effects as displayed in Figure 4. This gives the interesting result that the average annual M3 growth rates without and with adjustment between 2000Q3 and 2003Q2 equal 7.2 percent and 5.0 percent, respectively, i.e., adjusted money growth does not give a strong signal of inflationary pressure.

Another widely used measure of excess liquidity is the money overhang defined as the difference between observed money balances and (estimated) long–run money demand. In Figure 5, we compare two measures of money overhang, the first one being estimated from specification S5, i.e., without stock market variables, and the second one from specification S5c, i.e., including $r_t^e - r_t^o$ and $v_t$. It turns out that neglecting the stock market influence leads to a strong money overhang since the end of 2001 peaking in 2003Q2 with nearly 8%. This number compares well with the results of other authors who neglect the stock market influence, e.g., Belke et al. (2004). However, money overhang does not show up if the money demand function includes stock market variables. From this perspective, it can again be concluded that there are no serious inflationary threats of the recent excessive M3 growth rates once the stock market developments are taken into account.

## 6 Conclusion

In this paper the stability properties of various money demand specifications proposed in the literature have been analyzed. All specifications have in common that real money demand depends on income, interest rates and/or inflation. Using cointegration–breakdown tests recently introduced by Andrews and Kim (2003), the hypothesis of long–run structural stability has to be rejected for these specifications. The tests indicate that the break point is probably at the end of the year 2001 when M3 growth increased and stock market conditions deteriorated.

In an effort to restore a stable relationship between money and prices, the conventional money demand functions are augmented with variables from the financial sector: equity returns, real stock prices and stock market volatility. It turns out that these augmented specifications exhibit much better stability properties than the conventional ones. In particular, a specification including equity returns and stock market volatility passes all test. Moreover, short–run stability of single–equation error–correction models, which are estimated conditional on the long–run parameters, cannot be rejected.

To assess the importance of the stock market developments since the all-time high in 2000Q3, we simulate money demand conditional on unchanged and actually observed stock market variables and compare the two simulations with each other. The results suggests that the major part of M3 growth rates above the reference value of 4.5 percent is attributable to adverse stock market developments. This implies that inflationary pressure should be low.

The absence of inflationary pressure is reaffirmed by computing money overhang from the long–run money demand functions. It turns out that a conventional measure, which neglects the influence of the stock market variables, indicates a alarmingly high overhang while a more appropriate measure, which incorporates the influence of the stock market variables, does not indicate any noteworthy excess liquidity. This result strengthens the view put forward by the ECB that the actual money growth does not put any exceptional threat on price stability.

Overall, the results show that measures of excess liquidity come to different conclusions depending on their inclusion of stock market variables. Since the official target of 4.5 percent actual money growth is derived from a reasoning without stock market influences, it is no wonder that it does not work in the current situation. It was perhaps also due to this problem that the ECB decided to downweight the monetary pillar in their May 2003 policy revision.

## References

Andrews, D.W.K., 2003. End–of–sample instability tests. Econometrica 71, 1661-1694.

Andrews, D.W.K., and J.-Y. Kim, 2003. End–of–sample cointegration breakdown tests, Cowles Foundation Discussion Paper No. 1404.

Belke, A., W. Kösters, M. Leschke and T. Polleit, 20034. Challenges to ECB credibility, ECB Observer No. 5.

Belke, A., W. Kösters, M. Leschke and T. Polleit, 2004. Liquidity on the rise — too much money chasing too few goods, ECB Observer No. 6.

Brand, C., and N. Cassola, 2000. A money demand system for euro area M3, ECB Working Paper No. 39.

Brand, C., D. Gerdesmeier and B. Roffia, 2002. Estimating the trend of M3 income velocity underlying the reference value for monetary growth, ECB Occasional Paper No. 3.

Bruggeman, A., 2000. The stability of EMU-wide money demand functions and the monetary policy strategy of the European Central Bank. Manchester School 68, 184-202.

Bruggeman, A., P. Donati and A. Warne, 2003. Is the demand for Euro area M3 stable?, ECB Working Paper No. 255.

Calza, A., D. Gerdesmeier and J. Levy, 2001. Euro area money demand: Measuring the opportunity costs appropriately, IMF Working Paper No. 01/179.

Calza, A., and J. Sousa, 2003. Why has broad money demand been more stable in the euro area than in other economies? A literature review. In: O. Issing (ed.), Background studies for the ECB's evaluation of its monetary policy strategy, ECB, Frankfurt, 229-243.

Caruso, M., 2001. Stock prices and money velocity: a multi-country analysis. Empirical Economics 26, 651-672.

Carpenter, S.B., and J. Lange, 2002. Money demand and equity markets, FEDS Working Paper No. 2003-3.

Cassola, N., and C. Morana, 2002. Monetary policy and the stock market in the Euro area, ECB Working Papaer No. 119.

Choudhry, T., 1996. Real stock prices and the long–run money demand function: evidence from Canada and the USA. Journal of International Money and Finance 15, 1-17.

Clausen, V., and J.-R. Kim, 2000. The long–run stability of European money demand. Journal of Economic Integration 15, 486-505.

Coenen, G., and J.-L. Vega, 2001. The demand for M3 in the Euro area. Journal of Applied Econometrics 16, 727-748.

de Grauwe, P., 2003. The central bank that has missed the point, Financial Times, May 13.

ECB, 2003a. Economic developments in the euro area, ECB Monthly Bulletin, May.

ECB, 2003b. The ECB's monetary policy strategy, ECB Press Release, 8 May.

Elliott, G., 1999. Efficient tests for a unit root when the initial observation is drawn from its unconditional distribution. International Economic Review 40, 767-783.

Friedman, M., 1988. Money and the stock market. Journal of Political Economy 96, 221-245.

Funke, M., 2001. Money demand in Euroland. Journal of International Money and Finance 20, 701-713.

Golinelli, R., and S. Pastorello, 2002. Modelling the demand for M3 in the Euro Area. European Journal of Finance 8, 371-401.

Gottschalk, J., 1999. On the monetary transmission mechanism in Europe. Journal of Economics and Statistics 219, 357-374.

Hansen, B.E., 1992a. Testing for Parameter Instability in Linear Models. Journal of Policy Modeling 14, 517-533.

Hansen, B.E., 1992b. Tests for Parameter Instability in Regressions with I(1) Processes. Journal of Business and Economic Statistics 10, 321-335.

Hansen, H., and S. Johansen, 1999. Some tests for parameter constancy in cointegrated VAR-models. Econometrics Journal 2, 306-333.

Hayo, B., 1999. Estimating a European demand for money. Scottish Journal of Political Economy 46, 221-244.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 231-255.

Johansen, S., 1991. Estimation and hypothesis testing of cointegrating vectors in Gaussian vector autoregression models. Econometrica 59, 1551-1580.

Johansen, S., 2002. A small sample correction of the test for cointegration rank in the vector autoregressive model. Econometrica 70, 1929-62.

Kontolemis Z.G., 2002. Money demand in the Euro area: Where do we stand (today)?, IMF Working Paper No. 02/185.

Lütkepohl, H., and J. Wolters, 2003. The transmission of German monetary policy in the pre-Euro period. Macroeconomic Dynamics 7, forthcoming.

Müller, C., and E. Hahn, 2001. Money demand in Europe: Evidence from the past. Kredit und Kapital 34, 48-75.

Phillips, P.C.B., and B.E. Hansen, 1990. Statistical inference in instrumental variables regression with I(1) processes. Review of Economic Studies 57, 99-125.

Ploberger, W., W. Krämer and K. Kontrus, 1989. A new test for structural stability in the linear regression model. Journal of Econometrics 40, 307-318.

Stock, J.H., and M.W. Watson, 1993. A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 61, 783-820.

Wolters, J., T. Teräsvirta and H. Lütkepohl, 1998. Modelling the demand for M3 in the unified Germany. Review of Economics and Statistics 80, 399-409.

## A Construction of the data

The data set published by Calza et al. (2001) which contains data for M3, GDP, the GDP deflator, the long–term and the short–term interest rate, and the own rate of M3 from 1980Q1 until 1999Q4, is updated until 2003Q2. In order not to induce a break in the data series, we try to closely mimic their construction of variables. M3 is updated with flows adjusted for any changes which do not arise from transactions. This implies in particular that the break induced by EMU enlargement with the begin of 2001 is taken out of the data. In a similar manner, we update GDP and its price deflator by adding log changes to the last observation of the Calza et al. (2001) data set. Again, the EMU enlargement break is calculated out by using log changes of EMU-11 and EMU-12 before and after the enlargement, respectively. The short–term and long–term interest rates are updated with the 3–month money market rate and the 10–year government bond yield, respectively. Finally, the own rate of M3 is constructed from the rates of return to the components of M3 as outlined in Calza et al. (2001). The data for M3 and the interest rates from 2000Q1 until 2003Q2 are taken from the ECB homepage, the data for GDP and the GDP deflator are published by Eurostat and downloaded via Datastream.

Nominal stock prices are approximated by the German DAX30 from 1980 to 1986, because no European stock price index is available for this period, and the Dow Jones Euro Stoxx50 from 1987 to 2003. The data are downloaded from Datastream. The DAX30 is rescaled such that the value on 31 December 1986 equals the value of the Euro Stoxx50 on 1 January 1987. Quarterly nominal stock prices are constructed as quarterly averages of daily data obtained from Datastream. Dividing by the GDP deflator and taking logs yields real stock prices $sp_t$. Equity returns $r_t^e$ are constructed as the annualized three-year log differences of quarterly nominal stock prices. We use this rather long-term yield measure to mimic the fundamental yield path and exclude erratic short-term yield changes which probably do not affect long-run money demand. Stock market volatility is constructed as the two-year average of the conditional variance estimated from a leverage GARCH model with $t$–distributed innovations applied to daily yields of the nominal stock price index. Using averages makes the volatility index less erratic and better reveals the underlying movement in risk perception which, again, seems a better measure to include in a long-run money demand function.

Finally, we use the US dollar market price for UK Brent crude oil reported by the IMF International Financial Statistics and downloaded via Datastream. The price series is converted to Euro per barrel using the quarterly average of the US dollar exchange rate to the ECU/Euro.

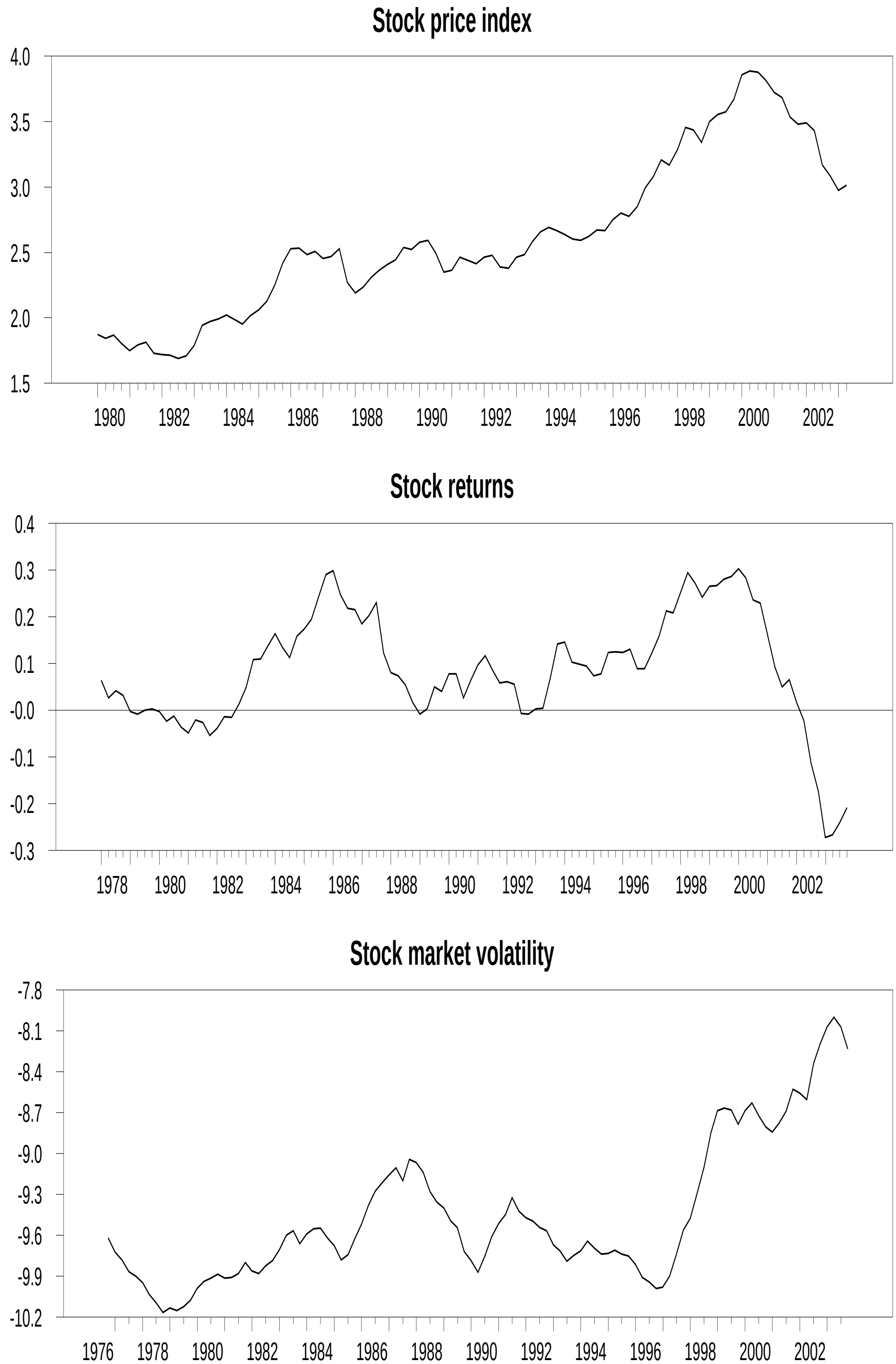



Fig. 1: The stock market variables from 1980Q1 to 2003Q2

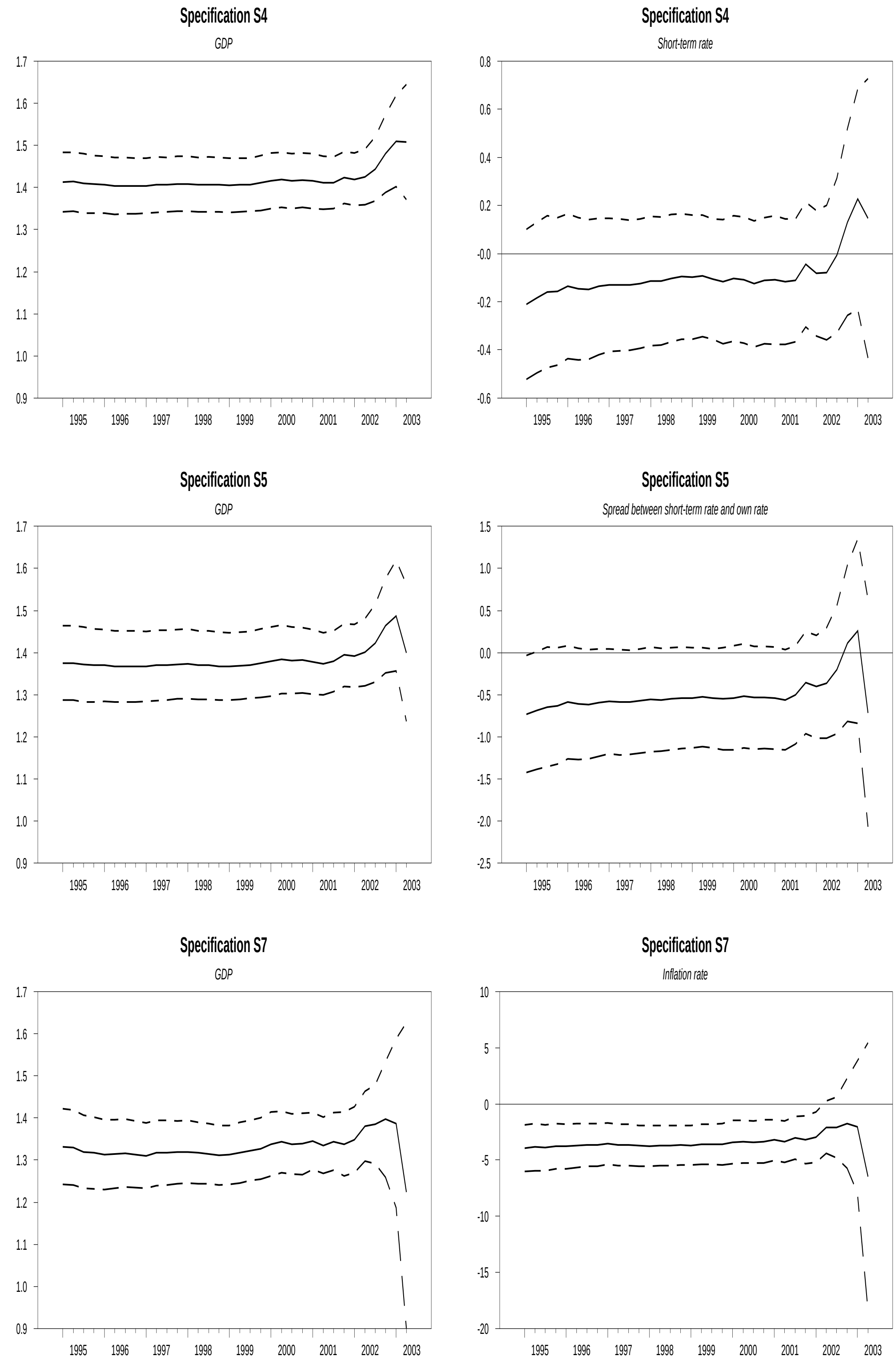



Fig. 2: Recursive FM–OLS estimates of the income elasticity (left panel) and the opportunity cost measure (right panel) for 1995Q1 to 2003Q2

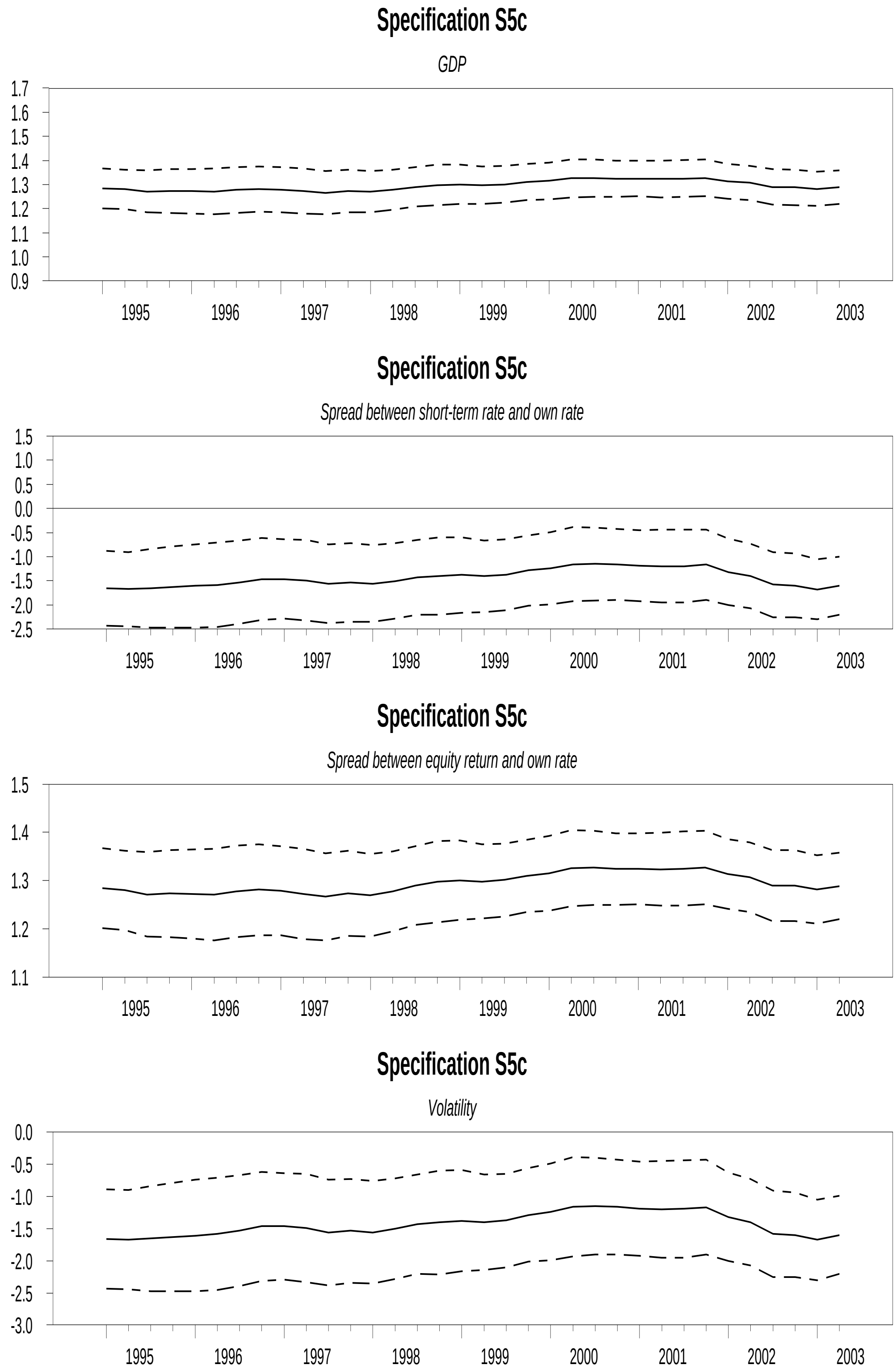
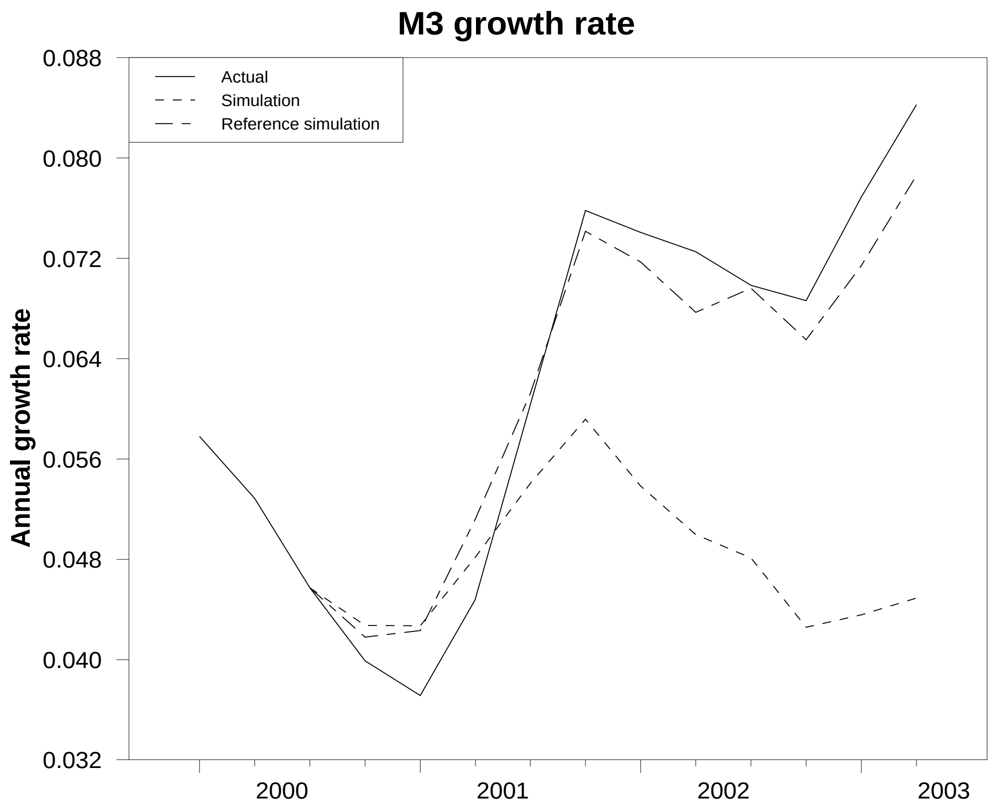



Fig. 3: Recursive FM–OLS estimates of specification S5c for 1995Q1 to 2003Q2



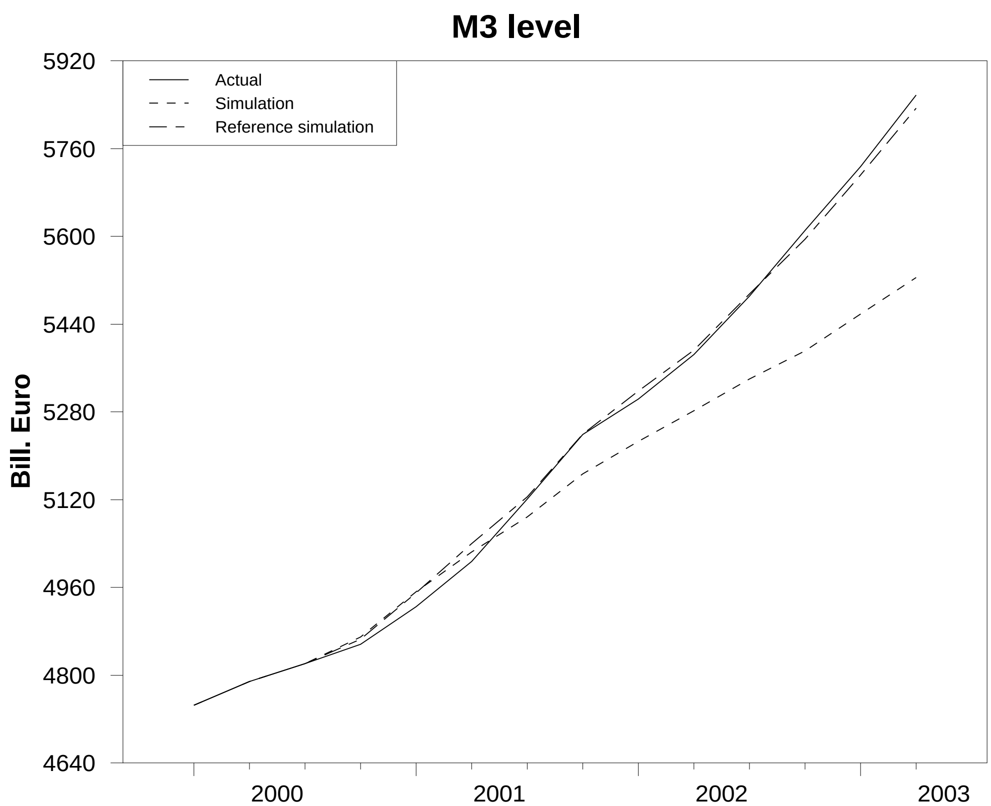




Fig. 4: Actual and simulated M3 for 2000Q1 to 2003Q2

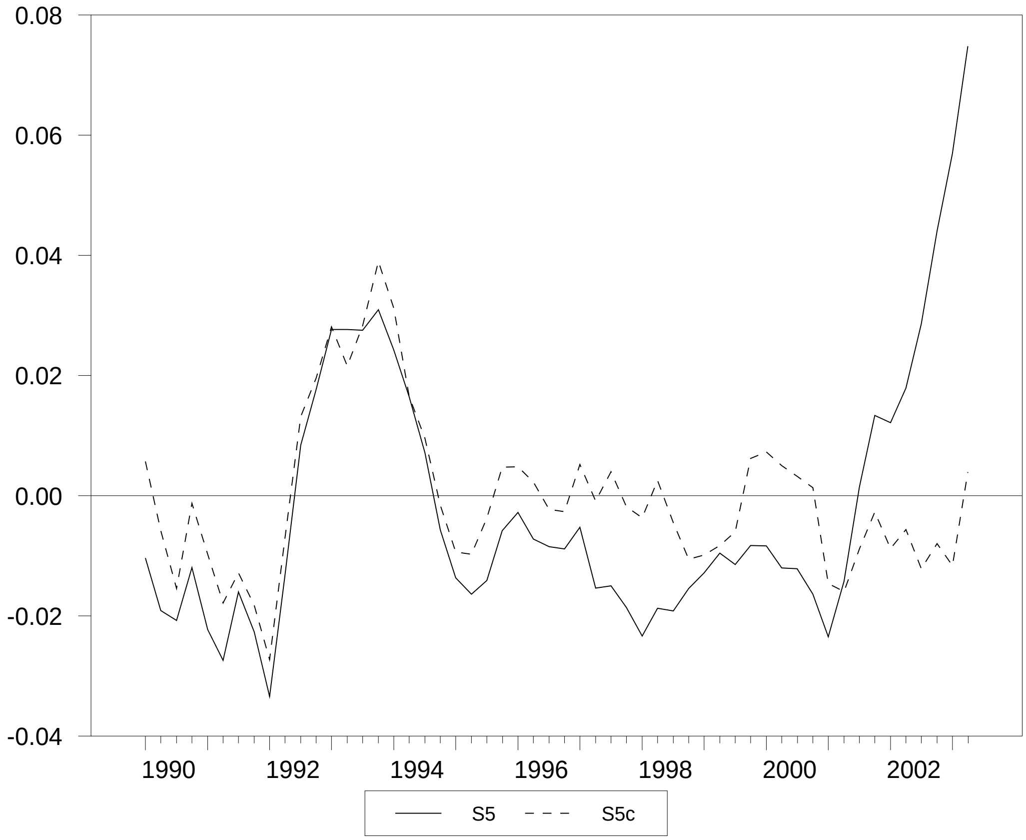



Fig. 5: Money overhang from 1990Q1 to 2003Q2

Tab. 1: Money demand specifications in the literature

| Specification | Restrictions | Authors |
|---|---|---|
| S1 | $\beta_3 = \beta_4 = \beta_5 = 0$ | Hayo (1999), Bruggeman (2000), Golinelli and Pastorello (2002) |
| S2 | $\beta_3 = \beta_5 = 0,\ \beta_2 = -\beta_4$ | Gottschalk (1999), Clausen and Kim (2000), Müller and Hahn (2001) |
| S3 | $\beta_3 = 0,\ \beta_2 = -\beta_4$ | Coenen and Vega (2001) |
| S4 | $\beta_2 = \beta_4 = \beta_5 = 0$ | Brand and Cassola (2000), Funke (2001), Kontolemis (2002) |
| S5 | $\beta_2 = \beta_5 = 0,\ \beta_3 = -\beta_4$ | Calza et al. (2001), Bruggeman et al. (2003) |
| S6 | $\beta_2 = 0,\ \beta_3 = -\beta_4$ | – |
| S7 | $\beta_2 = \beta_3 = \beta_4 = 0$ | Wolters et al. (1998), Lütkepohl and Wolters (2003) |
| S8 | $\beta_2 = \beta_3 = \beta_4 = \beta_5 = 0$ | – |

Notes: Only contributions published in the year 1999 and later are considered. For references to earlier contributions see Golinelli and Pastorello (2002).

Tab. 2: Cointegration of the conventional long-run money demand specifications in the baseline sample

| Specification | Bartlett corrected trace statistics | | | |
|---|---|---|---|---|
| | 4 | 3 | 2 | 1 |
| S1 | | 28.40<br>(0.072) | 8.98<br>(0.367) | 0.09<br>(0.771) |
| S2 | | 23.71<br>(0.214) | 8.90<br>(0.375) | 0.04<br>(0.850) |
| S3 | 44.99<br>(0.091) | 16.35<br>(0.688) | 6.75<br>(0.607) | 0.11<br>(0.738) |
| S4 | | 32.51<br>(0.024) | 9.44<br>(0.327) | 0.00<br>(0.971) |
| S5 | | 29.09<br>(0.060) | 9.51<br>(0.320) | 0.01<br>(0.904) |
| S6 | 62.90<br>(0.001) | 26.62<br>(0.111) | 8.01<br>(0.465) | 0.11<br>(0.739) |
| S7 | | 30.12<br>(0.046) | 7.03<br>(0.574) | 0.09<br>(0.761) |
| S8 | | | 11.11<br>(0.205) | 0.03<br>(0.872) |

Notes: The Bartlett corrected trace statistics proposed by Johansen (2002) are obtained from a VAR model with two lags which are sufficient to guarantee uncorrelated disturbances. The asymptotic $p$-values of the trace tests are given in brackets below. The computations are performed using Anders Warne's program Structural VAR 0.24.

Tab. 3: Estimates of the conventional money demand specifications for the baseline sample

| Specification | Estimation method | Estimated parameters | | | | |
|---|---|---|---|---|---|---|
| | | $\beta_1$ | $\beta_2$ | $\beta_3$ | $\beta_4$ | $\beta_5$ |
| S1 | OLS | 1.36 (0.026) | −0.32 (0.124) | | | |
| | FM–OLS | 1.34 (0.050) | −0.52 (0.237) | | | |
| | FIML | 1.37 (0.042) | −0.21 (0.221) | | | |
| S2 | OLS | 1.37 (0.028) | −0.51 (0.233) | | 0.51 (0.233) | |
| | FM–OLS | 1.39 (0.063) | −0.43 (0.538) | | 0.43 (0.538) | |
| | FIML | 1.55 (0.070) | 1.32 (0.613) | | −1.32 (0.613) | |
| S3 | OLS | 1.34 (0.029) | −0.34 (0.237) | | 0.34 (0.237) | −1.43 (0.606) |
| | FM–OLS | 1.33 (0.053) | 0.05 (0.422) | | −0.05 (0.422) | −3.40 (1.082) |
| | FIML | 1.35 (0.046) | 1.01 (0.364) | | −1.01 (0.364) | −5.48 (1.051) |
| S4 | OLS | 1.37 (0.022) | | −0.26 (0.088) | | |
| | FM–OLS | 1.39 (0.038) | | −0.20 (0.156) | | |
| | FIML | 1.37 (0.029) | | −0.28 (0.126) | | |
| S5 | OLS | 1.35 (0.023) | | −0.69 (0.170) | 0.69 (0.170) | |
| | FM–OLS | 1.35 (0.046) | | −0.74 (0.347) | 0.74 (0.347) | |
| | FIML | 1.32 (0.036) | | −0.94 (0.279) | 0.94 (0.279) | |
| S6 | OLS | 1.34 (0.024) | | −0.63 (0.230) | 0.63 (0.230) | −0.34 (0.747) |
| | FM–OLS | 1.34 (0.046) | | −0.45 (0.434) | 0.45 (0.434) | −1.29 (1.412) |
| | FIML | 1.25 (0.065) | | 4.72 (0.779) | −4.72 (0.779) | 20.64 (2.937) |
| S7 | OLS | 1.36 (0.024) | | | | −1.69 (0.584) |
| | FM–OLS | 1.31 (0.042) | | | | −3.70 (1.035) |
| | FIML | 1.27 (0.039) | | | | −4.73 (1.018) |
| S8 | OLS | 1.42 (0.016) | | | | |
| | FM–OLS | 1.43 (0.039) | | | | |
| | FIML | 1.42 (0.034) | | | | |

Notes: OLS denotes the ordinary least squares. FM–OLS denotes fully modified least squares. The non–parametric correction for FM–OLS is calculated using a Parzen kernel with associated automatic bandwidth selection as proposed by Hansen (1992b). FIML denotes the full information maximum likelihood estimator (Johansen, 1988, 1991) where we choose lag length 2 to obtain uncorrelated residuals and impose cointegration rank 1. Standard errors are reported in brackets below the estimates.

Tab. 4: Stability tests of the conventional money demand specifications for the baseline sample

| Test | S1 | S2 | S3 | S4 | S5 | S6 | S7 | S8 |
|---|---|---|---|---|---|---|---|---|
| $L_c$ | 0.53 (0.081) | 0.38 (0.189) | 0.41 (0.188) | 0.16 (0.540) | 0.25 (0.367) | 0.41 (0.185) | 0.39 (0.178) | 0.23 (0.337) |
| MeanQ | 1.14 (0.458) | 152.81 (0.044) | 0.74 (0.801) | 0.56 (0.351) | 0.45 (0.474) | 1.07 (0.401) | 0.34 (0.761) | 0.03 (0.976) |
| SupQ | 16.78 (0.421) | 3031.05 (0.052) | 2.88 (0.779) | 2.15 (0.347) | 2.46 (0.344) | 4.70 (0.331) | 2.11 (0.585) | 0.25 (0.884) |

Notes: $L_c$ denotes the Nyblom test based on FM–OLS estimation proposed by Hansen (1992b) who also provides surface response coefficients to approximate the asymptotic $p$–values which are given below the test statistics. MeanQ and SupQ denote the Nyblom tests proposed by Hansen and Johansen (1999), bootstrapped $p$–values calculated with Anders Warne's program Structural VAR 0.24 are given below the test statistics.

Tab. 5: Cointegration breakdown tests for the conventional long-run money demand specifications

| Test | S1 | S2 | S3 | S4 | S5 | S6 | S7 | S8 |
|---|---|---|---|---|---|---|---|---|
| *Break at 1999Q1, sample until 2001Q3* | | | | | | | | |
| $P_c$(OLS) | 0.935 | 0.952 | 0.984 | 0.903 | 0.887 | 0.887 | 0.919 | 0.855 |
| $R_c$(OLS) | 0.999 | 0.999 | 0.952 | 0.968 | 0.936 | 0.903 | 0.919 | 0.999 |
| $P_c$(FM–OLS) | 0.937 | 0.857 | 0.999 | 0.921 | 0.968 | 0.999 | 0.999 | 0.794 |
| $R_c$(FM–OLS) | 0.937 | 0.952 | 0.999 | 0.952 | 0.999 | 0.999 | 0.952 | 0.762 |
| $P_c$(FIML) | 0.871 | 0.823 | 0.936 | 0.903 | 0.952 | 0.000 | 0.952 | 0.774 |
| $R_c$(FIML) | 0.919 | 0.597 | 0.952 | 0.871 | 0.999 | 0.000 | 0.952 | 0.710 |
| *Break at 1999Q1, sample until 2003Q2* | | | | | | | | |
| $P_c$(OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| $R_c$(OLS) | 0.055 | 0.036 | 0.000 | 0.073 | 0.000 | 0.000 | 0.000 | 0.055 |
| $P_c$(FM–OLS) | 0.000 | 0.000 | 0.000 | 0.018 | 0.018 | 0.000 | 0.000 | 0.000 |
| $R_c$(FM–OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.018 | 0.750 | 0.000 | 0.071 |
| $P_c$(FIML) | 0.000 | 0.164 | 0.000 | 0.073 | 0.000 | 0.999 | 0.000 | 0.000 |
| $R_c$(FIML) | 0.273 | 0.564 | 0.000 | 0.091 | 0.000 | 0.999 | 0.000 | 0.473 |
| *Break at 2001Q4, sample until 2003Q2* | | | | | | | | |
| $P_c$(OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| $R_c$(OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| $P_c$(FM–OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.013 | 0.000 | 0.000 |
| $R_c$(FM–OLS) | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.180 | 0.000 | 0.000 |
| $P_c$(FIML) | 0.013 | 0.143 | 0.000 | 0.000 | 0.000 | 0.974 | 0.000 | 0.026 |
| $R_c$(FIML) | 0.000 | 0.156 | 0.000 | 0.000 | 0.000 | 0.974 | 0.000 | 0.039 |

Notes: $P_c$ and $R_c$ are the cointegration breakdown tests proposed by Andrews and Kim (2003). Only the simulated $p$–values are reported because the simulated critical values change from case to case so that the test statistics alone are difficult to interpret.

Tab. 6: Unit root tests for the stock market variables

| Test | Optimal lag length | | Lags for autocorrelation correction | | | | |
|---|---|---|---|---|---|---|---|
| | BIC | LM test | 1 | 2 | 3 | 4 | 12 |
| *Volatility* $v_t$ | | | | | | | |
| ADF | 1 | 1 | -0.33 | -0.48 | -0.44 | -0.48 | 0.18 |
| PP | | | 0.52 | 0.23 | 0.06 | -0.05 | 0.24 |
| DFGLSu | | | -0.62 | -0.79 | -0.74 | -0.91 | -0.35 |
| *Stock return* $r_t^e$ | | | | | | | |
| ADF | 12 | 12 | -1.29 | -1.30 | -1.87 | -1.97 | -1.97 |
| PP | | | -0.46 | -0.67 | -0.87 | -1.01 | -1.58 |
| DFGLSu | | | -1.41 | -1.34 | -1.96 | -1.93 | -1.61 |
| *Stock return spread* $r_t^e - r_t^o$ | | | | | | | |
| ADF | 1 | 12 | -1.54 | -1.46 | -2.14 | -2.12 | -1.93 |
| PP | | | -0.74 | -0.94 | -1.12 | -1.24 | -1.73 |
| DFGLSu | | | -1.57 | -1.51 | -2.10 | -2.06 | -1.71 |
| *Real stock prices* $sp_t$ | | | | | | | |
| ADF | 1 | 1 | -2.02 | -1.86 | -2.50 | -2.50 | -2.68 |
| PP | | | -1.37 | -1.53 | -1.68 | -1.81 | -2.02 |
| DFGLSu | | | -2.14 | -1.98 | -2.60 | -2.61 | -2.78 |

Notes: ADF is the augmented Dickey–Fuller test, PP is the Phillips–Perron test, and DFGLSu is the Dickey–Fuller test with GLS detrending proposed by Elliott (1999). The test regressions for $v_t$, $r_t^e$ and $r_t^e - r_t^o$ are estimated with a constant. The corresponding 5% critical values are -2.89 for ADF and PP, and -2.73 for DFGLSu. The test regressions for $sp_t$ are estimated with a constant and a trend. The corresponding 5% critical values are -3.45 for ADF and PP, and -3.17 for DFGLSu. The optimal lag length of the ADF model is determined both by the Bayesian information criterion (BIC) and by successively increasing the lag length until the null of no autocorrelation cannot be rejected using an LM test (LM).

Tab. 7: FM–OLS estimates of the augmented money demand specifications in the full sample

| Specification | Estimated parameters | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | $\beta_1$ | $\beta_2$ | $\beta_3$ | $\beta_4$ | $\beta_5$ | $\beta_6$ | $\beta_7$ | $\beta_8$ |
| S1a | 1.25 (0.062) | −1.30 (0.309) | | | | −0.12 (0.036) | | |
| S2a | 1.43 (0.124) | −0.85 (1.236) | | 0.85 (1.236) | | −0.09 (0.077) | | |
| S3a | 1.33 (0.117) | −1.34 (1.082) | | 1.34 (1.082) | −1.75 (2.983) | −0.05 (0.074) | | |
| S4a | 1.35 (0.038) | | −0.66 (0.167) | | | −0.13 (0.027) | | |
| S5a | 1.30 (0.047) | | −1.82 (0.418) | 1.82 (0.418) | | −0.14 (0.032) | | |
| S6a | 1.31 (0.048) | | −2.08 (0.477) | 2.08 (0.477) | 1.59 (1.452) | −0.13 (0.033) | | |
| S7a | 1.43 (0.086) | | | | −2.72 (2.783) | −0.06 (0.071) | | |
| S8a | 1.51 (0.062) | | | | | −0.12 (0.073) | | |
| S1b | 1.24 (0.091) | −1.27 (0.338) | | | | −0.12 (0.050) | 0.01 (0.025) | |
| S2b | 1.26 (0.163) | −0.53 (0.822) | | 0.53 (0.822) | | −0.16 (0.083) | 0.05 (0.039) | |
| S3b | 1.19 (0.126) | −0.46 (0.595) | | 0.46 (0.595) | −2.90 (1.677) | −0.17 (0.065) | 0.05 (0.028) | |
| S4b | 1.29 (0.068) | | −0.56 (0.176) | | | −0.16 (0.038) | 0.02 (0.018) | |
| S5b | 1.15 (0.076) | | −1.80 (0.342) | 1.80 (0.342) | | −0.22 (0.040) | 0.05 (0.018) | |
| S6b | 1.14 (0.080) | | −1.65 (0.391) | 1.65 (0.391) | −0.58 (1.215) | −0.22 (0.043) | 0.05 (0.018) | |
| S7b | 1.23 (0.123) | | | | −2.88 (1.703) | −0.17 (0.067) | 0.05 (0.029) | |
| S8b | 1.33 (0.153) | | | | | −0.17 (0.088) | 0.04 (0.042) | |
| S1c | 1.26 (0.059) | −1.08 (0.337) | | | | −0.10 (0.036) | | 0.013 (0.0104) |
| S2c | 1.37 (0.073) | −0.80 (0.702) | | 0.80 (0.702) | | −0.07 (0.044) | | 0.026 (0.0133) |
| S3c | 1.30 (0.058) | −0.39 (0.513) | | 0.39 (0.513) | −4.30 (1.415) | −0.11 (0.035) | | 0.023 (0.0095) |
| S4c | 1.34 (0.035) | | −0.55 (0.170) | | | −0.11 (0.026) | | 0.016 (0.0074) |
| S5c | 1.29 (0.035) | | −1.60 (0.308) | 1.60 (0.308) | | −0.12 (0.024) | | 0.021 (0.0065) |
| S6c | 1.29 (0.037) | | −1.49 (0.358) | 1.49 (0.358) | −0.31 (1.079) | −0.12 (0.025) | | 0.021 (0.0065) |
| S7c | 1.34 (0.047) | | | | −4.41 (1.401) | −0.12 (0.036) | | 0.022 (0.0096) |
| S8c | 1.45 (0.046) | | | | | −0.07 (0.045) | | 0.025 (0.0140) |

Notes: All specifications are estimated by FM–OLS with automatic bandwidth selection. The parameter $\beta_6$ measures the influence of $r_t^e$ in specifications S1 and S4, and the influence of $r_t^e - r_t^o$ otherwise. The parameters $\beta_7$ and $\beta_8$ measure the influence of real stock prices $sp_t$ and stock market volatility $v_t$, respectively. Standard errors are reported in brackets below the estimates.

Tab. 8: FIML estimates of the augmented money demand specifications in the full sample

| Specification | Estimated parameters | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | $\beta_1$ | $\beta_2$ | $\beta_3$ | $\beta_4$ | $\beta_5$ | $\beta_6$ | $\beta_7$ | $\beta_8$ |
| S1a | 1.13 (0.054) | −2.01 (0.306) | | | | −0.21 (0.040) | | |
| S2a | 2.38 (0.253) | 11.7 (2.634) | | −11.7 (2.634) | | 0.51 (0.184) | | |
| S3a | 1.49 (0.116) | 4.34 (1.047) | | −4.34 (1.047) | −15.1 (3.453) | −0.08 (0.086) | | |
| S4a | 1.29 (0.029) | | −0.97 (0.140) | | | −0.15 (0.026) | | |
| S5a | 1.22 (0.038) | | −2.58 (0.365) | 2.58 (0.365) | | −0.16 (0.032) | | |
| S6a | 1.28 (0.044) | | −3.49 (0.483) | 3.49 (0.483) | 6.04 (1.719) | −0.12 (0.036) | | |
| S7a | 1.02 (0.111) | | | | −18.8 (3.950) | −0.31 (0.100) | | |
| S8a | 1.63 (0.133) | | | | | −0.61 (0.172) | | |
| S1b | 1.10 (0.075) | −1.91 (0.321) | | | | −0.24 (0.049) | 0.02 (0.021) | |
| S2b | 5.75 (1.393) | 34.4 (7.421) | | −34.4 (7.421) | | 2.47 (0.837) | −0.52 (0.351) | |
| S3b | 0.80 (0.109) | 0.90 (0.456) | | −0.90 (0.456) | −12.6 (1.686) | −0.47 (0.067) | 0.14 (0.024) | |
| S4b | 1.23 (0.048) | | −0.84 (0.140) | | | −0.19 (0.033) | 0.02 (0.013) | |
| S5b | 1.15 (0.076) | | −1.80 (0.342) | 1.80 (0.342) | | −0.22 (0.040) | 0.05 (0.018) | |
| S6b | 1.14 (0.080) | | −1.65 (0.391) | 1.65 (0.391) | −0.58 (1.215) | −0.22 (0.043) | 0.05 (0.018) | |
| S7b | 1.23 (0.123) | | | | −2.88 (1.703) | −0.17 (0.067) | 0.05 (0.029) | |
| S8b | 1.33 (0.153) | | | | | −0.17 (0.088) | 0.04 (0.042) | |
| S1c | 1.20 (0.044) | −1.39 (0.255) | | | | −0.16 (0.032) | | 0.038 (0.0086) |
| S2c | 1.38 (0.068) | −0.25 (0.668) | | 0.25 (0.668) | | −0.06 (0.050) | | 0.084 (0.0158) |
| S3c | 1.27 (0.032) | 0.56 (0.261) | | −0.56 (0.261) | −7.77 (0.922) | −0.20 (0.025) | | 0.060 (0.0064) |
| S4c | 1.29 (0.023) | | −0.78 (0.114) | | | −0.13 (0.021) | | 0.030 (0.0057) |
| S5c | 1.25 (0.024) | | −1.87 (0.218) | 1.87 (0.218) | | −0.14 (0.020) | | 0.038 (0.0055) |
| S6c | 1.24 (0.023) | | −0.71 (0.252) | 0.71 (0.252) | −4.68 (0.954) | −0.18 (0.020) | | 0.052 (0.0053) |
| S7c | 1.24 (0.028) | | | | −7.23 (0.908) | −0.20 (0.025) | | 0.062 (0.0064) |
| S8c | 1.41 (0.038) | | | | | −0.06 (0.047) | | 0.083 (0.0155) |

Notes: All specifications are estimated by FIML with lag length 2 and cointegration rank 1. The parameter $\beta_6$ measures the influence of $r_t^e$ in specifications S1 and S4, and the influence of $r_t^e - r_t^o$ otherwise. The parameters $\beta_7$ and $\beta_8$ measure the influence of real stock prices $sp_t$ and stock market volatility $v_t$, respectively. Standard errors are reported in brackets below the estimates.

Tab. 9: Cointegration of the augmented long-run money demand specifications

| Specification | Bartlett corrected trace statistics | | | | | |
|---|---|---|---|---|---|---|
| | 6 | 5 | 4 | 3 | 2 | 1 |
| S1a | | | 44.80 (0.094) | 19.90 (0.430) | 6.56 (0.629) | 1.52 (0.218) |
| S2a | | | 34.64 (0.467) | 16.16 (0.701) | 4.86 (0.823) | 1.05 (0.306) |
| S3a | | 55.92 (0.381) | 30.23 (0.707) | 15.70 (0.734) | 4.72 (0.838) | 0.84 (0.359) |
| S4a | | | 47.43 (0.055) | 17.88 (0.575) | 4.49 (0.860) | 0.10 (0.755) |
| S5a | | | 42.29 (0.151) | 17.55 (0.599) | 5.50 (0.754) | 0.53 (0.466) |
| S6a | | 64.49 (0.124) | 35.64 (0.415) | 19.66 (0.446) | 5.98 (0.698) | 0.63 (0.428) |
| S7a | | | 35.96 (0.398) | 19.48 (0.457) | 5.72 (0.728) | 0.70 (0.403) |
| S8a | | | | 17.15 (0.629) | 5.41 (0.764) | 0.82 (0.366) |
| S1b | | 51.90 (0.554) | 26.11 (0.886) | 11.64 (0.944) | 4.92 (0.817) | 0.95 (0.329) |
| S2b | | 42.45 (0.901) | 23.52 (0.952) | 11.12 (0.958) | 3.88 (0.913) | 0.13 (0.714) |
| S3b | 82.40 (0.289) | 42.76 (0.894) | 22.76 (0.965) | 10.05 (0.980) | 4.02 (0.902) | 0.11 (0.737) |
| S4b | | 60.37 (0.224) | 28.35 (0.798) | 11.23 (0.955) | 4.12 (0.894) | 0.03 (0.860) |
| S5b | | 58.16 (0.296) | 26.29 (0.880) | 11.84 (0.937) | 4.35 (0.873) | 0.18 (0.675) |
| S6b | 95.19 (0.055) | 56.14 (0.372) | 27.39 (0.839) | 11.44 (0.950) | 5.00 (0.809) | 0.20 (0.655) |
| S7b | | 63.81 (0.137) | 25.94 (0.891) | 10.89 (0.964) | 4.99 (0.810) | 0.42 (0.518) |
| S8b | | | 25.17 (0.914) | 12.16 (0.926) | 4.78 (0.832) | 0.43 (0.511) |
| S1c | | 67.15 (0.080) | 29.59 (0.739) | 16.39 (0.684) | 4.14 (0.892) | 0.80 (0.372) |
| S2c | | 57.36 (0.325) | 30.33 (0.702) | 14.06 (0.837) | 3.12 (0.961) | 0.17 (0.684) |
| S3c | 103.07 (0.014) | 47.38 (0.747) | 24.21 (0.938) | 13.45 (0.870) | 3.05 (0.965) | 0.04 (0.847) |
| S4c | | 74.74 (0.019) | 29.62 (0.738) | 15.23 (0.765) | 2.57 (0.983) | 0.01 (0.923) |
| S5c | | 75.97 (0.015) | 31.79 (0.624) | 15.01 (0.779) | 2.88 (0.972) | 0.04 (0.836) |
| S6c | 118.20 (0.001) | 58.15 (0.297) | 28.73 (0.781) | 15.31 (0.760) | 3.21 (0.956) | 0.01 (0.908) |
| S7c | | 84.14 (0.002) | 28.82 (0.777) | 15.97 (0.714) | 2.70 (0.979) | 0.03 (0.853) |
| S8c | | | 40.36 (0.210) | 13.95 (0.844) | 2.83 (0.974) | 0.14 (0.713) |

Notes: The Bartlett corrected trace statistics proposed by Johansen (2002) are obtained from a VAR model with two lags which are sufficient to guarantee uncorrelated disturbances. The asymptotic $p$-values of the trace tests are given in brackets below. The computations are performed using Anders Warne's program Structural VAR 0.24.

Tab. 10: Cointegration breakdown tests for the augmented long-run money demand specifications in the full sample

| Test | S1a | S2a | S3a | S4a | S5a | S6a | S7a | S8a |
|---|---|---|---|---|---|---|---|---|
| *Break at 1999Q1, sample until 2001Q3* | | | | | | | | |
| $P_c$(FM–OLS) | 0.905 | 0.857 | 0.921 | 0.999 | 0.999 | 0.999 | 0.984 | 0.810 |
| $R_c$(FM–OLS) | 0.921 | 0.873 | 0.999 | 0.937 | 0.999 | 0.999 | 0.999 | 0.651 |
| *Break at 1999Q1, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.964 | 0.714 | 0.237 | 0.982 | 0.911 | 0.929 | 0.196 | 0.536 |
| $R_c$(FM–OLS) | 0.999 | 0.893 | 0.446 | 0.982 | 0.571 | 0.750 | 0.411 | 0.893 |
| *Break at 2001Q4, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.782 | 0.359 | 0.026 | 0.885 | 0.962 | 0.962 | 0.000 | 0.436 |
| $R_c$(FM–OLS) | 0.667 | 0.244 | 0.000 | 0.859 | 0.936 | 0.936 | 0.000 | 0.474 |
| Test | S1b | S2b | S3b | S4b | S5b | S6b | S7b | S8b |
| *Break at 1999Q1, sample until 2001Q3* | | | | | | | | |
| $P_c$(FM–OLS) | 0.889 | 0.841 | 0.921 | 0.968 | 0.999 | 0.999 | 0.921 | 0.809 |
| $R_c$(FM–OLS) | 0.857 | 0.921 | 0.857 | 0.746 | 0.921 | 0.937 | 0.794 | 0.730 |
| *Break at 1999Q1, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.946 | 0.589 | 0.536 | 0.964 | 0.964 | 0.964 | 0.375 | 0.482 |
| $R_c$(FM–OLS) | 0.999 | 0.875 | 0.927 | 0.946 | 0.927 | 0.927 | 0.839 | 0.768 |
| *Break at 2001Q4, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.756 | 0.282 | 0.167 | 0.885 | 0.936 | 0.936 | 0.141 | 0.244 |
| $R_c$(FM–OLS) | 0.654 | 0.167 | 0.090 | 0.936 | 0.949 | 0.974 | 0.090 | 0.205 |
| Test | S1c | S2c | S3c | S4c | S5c | S6c | S7c | S8c |
| *Break at 1999Q1, sample until 2001Q3* | | | | | | | | |
| $P_c$(FM–OLS) | 0.952 | 0.857 | 0.968 | 0.999 | 0.984 | 0.999 | 0.937 | 0.778 |
| $R_c$(FM–OLS) | 0.841 | 0.889 | 0.873 | 0.746 | 0.937 | 0.937 | 0.825 | 0.714 |
| *Break at 1999Q1, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.857 | 0.875 | 0.929 | 0.982 | 0.964 | 0.982 | 0.927 | 0.661 |
| $R_c$(FM–OLS) | 0.750 | 0.714 | 0.964 | 0.732 | 0.714 | 0.821 | 0.999 | 0.554 |
| *Break at 2001Q4, sample until 2003Q2* | | | | | | | | |
| $P_c$(FM–OLS) | 0.821 | 0.641 | 0.449 | 0.846 | 0.795 | 0.808 | 0.603 | 0.654 |
| $R_c$(FM–OLS) | 0.795 | 0.628 | 0.654 | 0.923 | 0.628 | 0.667 | 0.782 | 0.731 |

Notes: $P_c$ and $R_c$ are the cointegration breakdown tests proposed by Andrews and Kim (2003). Only the simulated $p$–values are reported because the simulated critical values change from case to case.

Tab. 11: Stability tests of the augmented money demand specifications for the full sample

| Test | S1c | S2c | S3c | S4c | S5c | S6c | S7c | S8c |
|---|---|---|---|---|---|---|---|---|
| $L_c$ | 0.76 | 0.61 | 0.62 | 0.14 | 0.21 | 0.23 | 0.29 | 0.32 |
| | (0.095) | (0.187) | (0.277) | (0.983) | (0.801) | (0.961) | (0.624) | (0.291) |
| MeanQ | 1.10 | 4.40 | 0.46 | 0.37 | 0.10 | 1.00 | 0.42 | 1.49 |
| | (0.210) | (0.169) | (0.651) | (0.567) | (0.969) | (0.311) | (0.484) | (0.228) |
| SupQ | 5.21 | 11.68 | 2.62 | 1.18 | 0.28 | 3.33 | 2.68 | 3.43 |
| | (0.093) | (0.196) | (0.249) | (0.465) | (0.969) | (0.198) | (0.108) | (0.261) |

Notes: $L_c$ denotes the Nyblom test based on FM–OLS estimation proposed by Hansen (1992b) who also provides surface response coefficients to approximate the asymptotic $p$–values which are given below the test statistics. MeanQ and SupQ denote the Nyblom tests proposed by Hansen and Johansen (1999) computed for the sample 1999Q1 to 2003Q2; bootstrapped $p$–values calculated with Anders Warne's program Structural VAR 0.24 are given below the test statistics.

Tab. 12: Tests for short–run stability of the augmented money demand specifications for the full sample

| Equation | S1c | S2c | S3c | S4c | S5c | S6c | S7c | S8c |
|---|---|---|---|---|---|---|---|---|
| *Ploberger et al. (1989) fluctuation test for VAR equations* | | | | | | | | |
| $mp_t$ | 0.92 (0.976) | 2.79 (0.058) | 1.92 (0.169) | 1.17 (0.806) | 1.29 (0.637) | 1.82 (0.215) | 2.08 (0.102) | 2.51 (0.070) |
| $y_t$ | 1.87 (0.246) | 1.82 (0.316) | 1.70 (0.289) | 1.85 (0.224) | 1.85 (0.164) | 1.74 (0.285) | 1.61 (0.319) | 1.70 (0.340) |
| $r^l_t$, $r^l_t - r^o_t$, $r^l_t$, $r^s_t - r^o_t$ | 1.08 (0.879) | 1.19 (0.816) | 1.10 (0.892) | 1.85 (0.212) | 1.55 (0.372) | 1.29 (0.729) | 1.00 (0.888) | |
| $\Delta p_t$ | | | 1.32 (0.671) | | | 1.17 (0.829) | | |
| $r^e_t$, $r^e_t - r^o_t$ | 1.11 (0.888) | 2.36 (0.104) | 1.14 (0.858) | 1.10 (0.873) | 1.11 (0.838) | 1.12 (0.859) | 1.10 (0.850) | 2.23 (0.100) |
| $v_t$ | 1.76 (0.319) | 1.76 (0.376) | 1.85 (0.225) | 1.82 (0.252) | 1.78 (0.236) | 1.90 (0.198) | 1.86 (0.217) | 1.78 (0.311) |
| *Hansen (1992) Nyblom test* | | | | | | | | |
| $mp_t$ | 2.68 (0.106) | 2.18 (0.326) | 2.75 (0.089) | 2.19 (0.324) | 2.19 (0.316) | 2.27 (0.271) | 2.61 (0.128) | 2.11 (0.377) |
| *Andrews (2003) stability tests for break at 1999Q1, sample until 2001Q3* | | | | | | | | |
| $mp_t$ | 0.279 | 0.426 | 0.311 | 0.721 | 0.885 | 0.869 | 0.361 | 0.492 |
| *Andrews (2003) stability tests for break at 1999Q1, sample until 2003Q2* | | | | | | | | |
| $mp_t$ | 0.889 | 0.963 | 0.981 | 0.611 | 0.537 | 0.574 | 0.999 | 0.981 |
| *Andrews (2003) stability tests for break at 2001Q4, sample until 2003Q2* | | | | | | | | |
| $mp_t$ | 0.447 | 0.408 | 0.368 | 0.632 | 0.842 | 0.829 | 0.408 | 0.382 |

Notes: The Ploberger et al. (1989) fluctuations tests of each VAR equation are given together with bootstrapped $p$–values computed with Anders Warne's program Structural VAR 0.24. The Hansen (1992) Nyblom test and the Andrews (2003) stability tests are applied to the single–equation error–correction models for $mp_t$. For the Nyblom tests, $p$–values are calculated from simulating the asymptotic distribution. For the Andrews stability tests, only $p$–values are reported; they are obtained from parametric subsampling.

Tab. 13: Decomposition of the estimated long–run money demand changes

| Year | $\Delta\hat{mp}^{lr}_t$ | Conventional | $\Delta y_t$ | $\Delta r^s_t$ | $\Delta r^o_t$ | Stock market | $\Delta r^e_t$ | $\Delta v_t$ |
|---|---|---|---|---|---|---|---|---|
| 1995 | 1.5 | 1.9 | 2.0 | -0.0 | -0.1 | -0.4 | -0.4 | -0.0 |
| 1996 | 3.5 | 3.5 | 2.2 | 3.1 | -1.7 | -0.1 | 0.4 | -0.5 |
| 1997 | 2.8 | 3.3 | 3.6 | 0.2 | -0.6 | -0.5 | -1.5 | 0.9 |
| 1998 | 4.2 | 3.0 | 2.5 | 1.3 | -0.8 | 1.1 | -0.4 | 1.5 |
| 1999 | 3.1 | 3.5 | 3.9 | 0.3 | -0.6 | -0.4 | -0.6 | 0.1 |
| 2000 | 2.8 | 2.2 | 3.5 | -2.5 | 1.2 | 0.7 | 0.7 | -0.0 |
| 2001 | 5.2 | 2.6 | 1.0 | 2.5 | -0.9 | 2.6 | 2.0 | 0.6 |
| 2002 | 5.3 | 1.6 | 1.4 | 0.5 | -0.3 | 3.6 | 2.9 | 0.7 |
| 2003 | 4.1 | 0.9 | -0.3 | 2.4 | -1.2 | 3.1 | 2.3 | 0.8 |

Notes: $\Delta\hat{mp}^{lr}_t$ is the estimated growth rate of long–run real money demand. For 1995 to 2002, annual growth rates are reported. For 2003, the annualized growth rate for the first half of the year is reported.