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**Kiel Working Paper No. 1179**

**Is European Money Demand Still Stable?**

by

**Kai Carstensen**

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# Is European money demand still stable?

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## **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 overshoot 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 real stock prices and volatility remain stable. Using such an augmented specification, the claim that the excessive M3 growth rates are due to adverse stock market developments is examined. The results indicate that one cannot expect these growth rates to revert in the near future unless one is willing to assume a quick recovery of the European stock markets.

## **1 Introduction**

On 8 May 2003 the ECB announced a revision of its monetary policy strategy (ECB, 2003b). While it continues to follow a two-pillar strategy, analysis of the monetary developments is now mainly intended to assess medium to long-term trends in inflation. Since money is not any longer assigned a “prominent role” in the conduct of monetary policy, the revision can be interpreted as a weakening of the first pillar. This step 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 3.25 percent at the end of August 2001 to 2 percent at the beginning of August 2003 although the monetary developments suggested opposite action.

The ECB explains the strong money growth with portfolio shifts from equities to safe and liquid assets which are induced by financial uncertainty and will be reversed once uncertainty diminishes (e.g. ECB, 2003a). From this perspective, the recent money growth does not 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 (e.g., Gottschalk, 1999, Hayo, 1999, Bruggeman, 2000, Clausen and Kim, 2000, Coenen and Vega, 2001, Funke, 2001, Müller and Hahn, 2001, Golinelli and Pastorello, 2002) or up to the first year of EMU (Brand and Cassola, 2000, Calza et al., 2001) 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. Moreover, the asymptotic distributions of the fluctuation and Nyblom-type tests might be poor approximations for the finite sample at hand. In particular, asymptotic distributions are generally derived under the assumption that the post-break sample size is a fixed fraction of the total sample size and, thus, tends to infinity with the full sample size. For small post-break sample sizes this might not be an adequate assumption.

This paper adds to the literature by, first, using an updated data set from the first quarter of 1980 until the fourth quarter of 2002. Consequently, there are more observations with excessive money growth at the end of the sample available. Second, we use 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. They assume a fixed and finite post-break sample size even asymptotically. Moreover, critical values and  $p$ -values can easily be obtained by parametric sub-sampling. In addition, we also test for short-run instability using a similar test put forward by Andrews (2002) for stationary environments. Since we find conventional money demand specifications to become unstable in 2001, we specify a money demand function augmented by some financial variables which exhibits structural stability and can be used to quantitatively assess the importance of stock market developments on M3 growth rates.

The remainder of this paper is organized as follows. In Section 2 we present a number of money demand specifications typically used to model EMU money demand. In Section 3 we outline the end-of-sample stability tests proposed by Andrews (2002) and Andrews and Kim (2003). The empirical test results and alternative specifications are described in Section 4 while some policy implications which follow from the quantitative importance of the stock market developments are discussed in Section 5. Finally, Section 5 concludes.

## 2 Specifications of the European money demand function

In this section we give 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  $\Delta p_t$ ,

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

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-

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)

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

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). Since we have a direct measure of the own rate at hand and since it is the spread between the two interest rates which should matter, we include GDP and the spread between the long-term rate and the own rate in 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. Finally, we use a variant where money demand solely depends on GDP and the inflation rate (specification S6). 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.

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 until 2002Q4. 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 2002Q4. In order not to induce a break in the data series, we try to closely mimic their construction of variables. We update M3 with the help of flows adjusted for any changes which do not arise from transactions. In particular, 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 in their data set. Again, an EMU enlargement break is calculated out. 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 2000Q1 until 2002Q4 are taken from the ECB Homepage. 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. 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 can 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 (2002) 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  $t = 1, \dots, T + m$  into the first  $T$  and the last  $m$  observations, they start from the linear model

$$y_t = \begin{cases} x_t' \beta_0 + u_t, & t = 1, \dots, T \\ x_t' \beta_t + u_t, & t = T + 1, \dots, T + m, \end{cases} \quad (2)$$

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  $t = T + 1, \dots, T + m$



and  $u_t$  is stationary for all  $t = 1, \dots, T + m$ , against the alternative hypothesis that either  $\beta_0 \neq \beta_t$  for some  $t \in \{T + 1, \dots, T + m\}$  or the distribution of  $\{u_{T+1}, \dots, u_{T+m}\}$  differs from the distribution of  $\{u_1, \dots, u_m\}$ . 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.

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

$$\hat{u}_t = y_t - x'_t \hat{\beta}_{1-T}, \quad t = T + 1, \dots, T + m, \quad (3)$$

from which the sum-of-squares statistic

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

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.<sup>1</sup>

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 the statistic

$$P_1(\beta_0) = \sum_{t=1}^m (y_t - x'_t \beta_0)^2$$

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

$$P_j(\beta_0) = \sum_{t=j}^{j+m-1} (y_t - x'_t \beta_0)^2, \quad j = 1, \dots, T - m + 1,$$

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<sup>1</sup> Note that  $[m/2]$  denotes the smallest integer greater than or equal to  $m/2$ .

are stationary and ergodic, the empirical distribution function of  $P_j(\beta_0)$ ,  $j = 1, \dots, T - m + 1$ , 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  $t = 1, \dots, T$  and the prediction sample  $t = T + 1, \dots, T + m$  are non-overlapping, Andrews and Kim suggest to evaluate  $P_j$  at the “leave- $m$ -out” estimator  $\hat{\beta}_{(j)}$  which uses the observations  $t = 1, \dots, T$  with  $t \neq j, \dots, j + m - 1$ . 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  $t = T + 1, \dots, T + m$ . A test statistic analogous to  $P_a$  is defined by the weighted sum

$$R_a = \sum_{i=T+1}^{T+m} \sum_{j=T+1}^{T+m} \min\{i - T, j - T\} \hat{u}_i \hat{u}_j = \sum_{i=T+1}^{T+m} \left( \sum_{j=i}^{T+m} \hat{u}_j \right)^2, \quad (5)$$

where  $\hat{u}_i$ ,  $i = T + 1, \dots, T + m$ , 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 (5) 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 alterna-

tive hypothesis of the disturbances changing from being stationary to being integrated  $I(1)$ . Surprisingly, the former test seems to possess more power against the latter hypothesis and vice versa. Thus, we will use the  $P_c$  statistic to test for stationary disturbances and the  $R_c$  statistic to test for parameter stability.

### 3.2 Stability tests in stationary regressions

The end-of-sample stability tests for stationary regressions proposed by Andrews (2002) 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 autocorrelated 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  $t = T + 1, \dots, T + m$  and the distribution of all  $u_i$ ,  $i = T + 1, \dots, T + m$ , equals the distribution of  $u_i$ ,  $i = 1, \dots, T$ , against the alternative hypothesis that either  $\beta_0 \neq \beta_t$  for some  $t \in \{T + 1, \dots, T + m\}$  or the distribution of some  $u_i$ ,  $i = T + 1, \dots, T + m$ , differs from the distribution of  $u_i$ ,  $i = 1, \dots, T$ .

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

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

where  $\hat{U}_t = (\hat{u}_t, \dots, \hat{u}_{t+m-1})'$  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  $V_t = \hat{\Sigma}^{-1/2} U_t$ ,  $\bar{Y}_t = \hat{\Sigma}^{-1/2} (y_t, \dots, y_{t+m-1})'$  and  $\bar{X}_t = \hat{\Sigma}^{-1/2} (x_t, \dots, x_{t+m-1})'$  then yields a model with i.i.d. disturbances

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

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

$$S_d = \hat{V}'_{T+1}\bar{X}'_{T+1}(\bar{X}'_{T+1}\bar{X}_{T+1})^{-1}\bar{X}_{T+1}\hat{V}_{T+1}, \quad (6)$$

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

$$S_d = \hat{V}'_{T+1}\hat{V}_{T+1}. \quad (7)$$

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.

## 4 Empirical test results

In this section we present and discuss the results of the stability tests applied to several specifications of the Euro area money demand function. We proceed as follows. First, we estimate each specification for the baseline sample 1980Q1 to 1998Q4 by means of OLS, FM-OLS, dynamic GLS (DGLS) as proposed by Stock and Watson (1993) and full information maximum likelihood (FIML) as proposed by Johansen (1988, 1991). We then graphically report recursive parameter estimates and, subsequently, cointegration-breakdown tests of Andrews and Kim (2003) for the quarters 1999Q1 to 2002Q4 in order to test for long-run stability in this period. Since we find long-run instability, we propose augmented specification to restore stability. Finally, following the suggestion by Hansen (1992a), we take the cointegration parameters as given, put up an error-correction model for each specification and test for short-run stability.

### 4.1 Baseline estimates

The estimated parameters of the seven specifications S1 to S7 of the money demand function (1) are displayed in Table 2. The income elasticity  $\beta_1$  is estimated remarkably stable as roughly 1.4 over specifications and estimation methods. This stability is also documented by Brand et al. (2002) who use data until 2001Q2. The semi-elasticity for the

Tab. 2: Estimates of the money demand specifications for the baseline sample

Specification	Estimation method	Estimated parameters				
		$\beta_1$	$\beta_2$	$\beta_3$	$\beta_4$	$\beta_5$
S1	OLS	1.39	−0.26			
	FM-OLS	1.42 (0.056)	−0.25 (0.271)			
	DGLS	1.40 (0.062)	−0.21 (0.333)			
	FIML	1.40 (0.046)	−0.10 (0.235)			
S2	OLS	1.41	−0.30		0.30	
	FM-OLS	1.49 (0.070)	0.24 (0.617)		−0.24 (0.617)	
	DGLS	1.42 (0.069)	−0.20 (0.645)		0.20 (0.645)	
	FIML	2.31 (0.236)	8.37 (2.123)		−8.37 (2.123)	
S3	OLS	1.36	−0.17		0.17	−1.77
	FM-OLS	1.37 (0.050)	0.63 (0.400)		−0.63 (0.400)	−4.65 (1.032)
	DGLS	1.37 (0.064)	0.34 (0.618)		−0.34 (0.618)	−3.11 (1.857)
	FIML	1.35 (0.049)	1.15 (0.381)		−1.15 (0.381)	−5.80 (1.114)
S4	OLS	1.39		−0.26		
	FM-OLS	1.41 (0.042)		−0.20 (0.176)		
	DGLS	1.40 (0.037)		−0.16 (0.174)		
	FIML	1.40 (0.032)		−0.15 (0.140)		
S5	OLS	1.36		−0.71	0.71	
	FM-OLS	1.37 (0.053)		−0.79 (0.404)	0.79 (0.404)	
	DGLS	1.36 (0.045)		−0.65 (0.362)	0.65 (0.362)	
	FIML	1.35 (0.041)		−0.69 (0.320)	0.69 (0.320)	
S6	OLS	1.35		−0.54	0.54	−0.75
	FM-OLS	1.35 (0.044)		−0.34 (0.412)	0.34 (0.412)	−1.89 (1.336)
	DGLS	1.35 (0.056)		−0.42 (0.632)	0.42 (0.632)	−1.09 (2.570)
	FIML	1.25 (0.068)		−5.07 (0.816)	5.07 (0.816)	21.96 (3.074)
S7	OLS	1.37				−1.88
	FM-OLS	1.32 (0.041)				−4.30 (1.000)
	DGLS	1.35 (0.064)				−2.48 (1.755)
	FIML	1.28 (0.039)				−4.61 (1.022)

Notes: OLS denotes the ordinary least squares estimator. FM-OLS denotes fully modified OLS with automatic bandwidth selection. DGLS denotes the dynamic generalized least squares estimator with two leads and lags in the dynamic equation as well as two lags for the covariance estimation. This is motivated by two reasons: First, Stock and Watson (1993) recommend this choice for a sample size of  $T = 100$  which is similar to ours. Second, two lags are sufficient in a VAR framework to obtain uncorrelated errors. FIML denotes the full information maximum likelihood (Johansen) estimator where we choose lag length 2 and impose cointegration rank 1. Standard errors are reported in brackets below

long-term rate in specification S1 has the expected negative sign but is estimated quite imprecisely. Using the spread between the long-term rate and the own rate instead makes things even worse: We obtain implausible and/or insignificant estimates both in specification S2 (without inflation rate) and specification S3 (with inflation rate).<sup>2</sup> In specification S4, the semi-elasticity for the short-term rate again has the expected sign but is estimated imprecisely. This changes if we use the spread between the short-term rate and the long-term rate instead. As one would expect, the estimated parameter is significantly smaller than zero. However, adding the inflation rate in specification S6 again asks too much of the data leading to insignificant estimates. We therefore drop all interest rates in specification S7 where money demand depends solely on GDP and the inflation rate, and obtain significant and plausible estimates. From the estimation results we conclude that specifications S5 and S7 are the most promising candidates for the subsequent stability analysis. In addition, we will also consider specifications S1 and S4 even if the interest rate parameters are estimated rather imprecisely.

## 4.2 Tests for long-run stability

To obtain a first impression of the stability properties of EMU money demand, we present recursive FM-OLS parameter estimates together with 90 percent confidence intervals for the specifications S1, S4, S5 and S7, see Figure 1. There is only little variation of the estimated parameters during the time from 1999Q1 to 2001Q4. Only at the end of the sample there is some sign of instability. However, it does not look too serious compared to the parameter shifts at the beginning of the nineties, which according to the results in the literature do not imply any structural break. A possible reason for this is the observation that recursive parameter estimates might be misleading as an indicator of structural instability simply because they are correlated with each other and over time. The lack of statistical guidance as to whether parameter shifts are really significant, is the

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<sup>2</sup> For specification S3 the results stand in contrast to the results presented by Coenen and Vega (2001) who used the long-run solution of an autoregressive distributed lag (ADL) model to estimate the cointegration parameters. However, taking their data set and employing OLS, FM-OLS and DGLS we obtain results similar to ours. Moreover, applying the ADL method to our data set, we obtain estimates similar to the ones presented in Table 2. Obviously, their specific data set and method play an important role to generate plausible results.

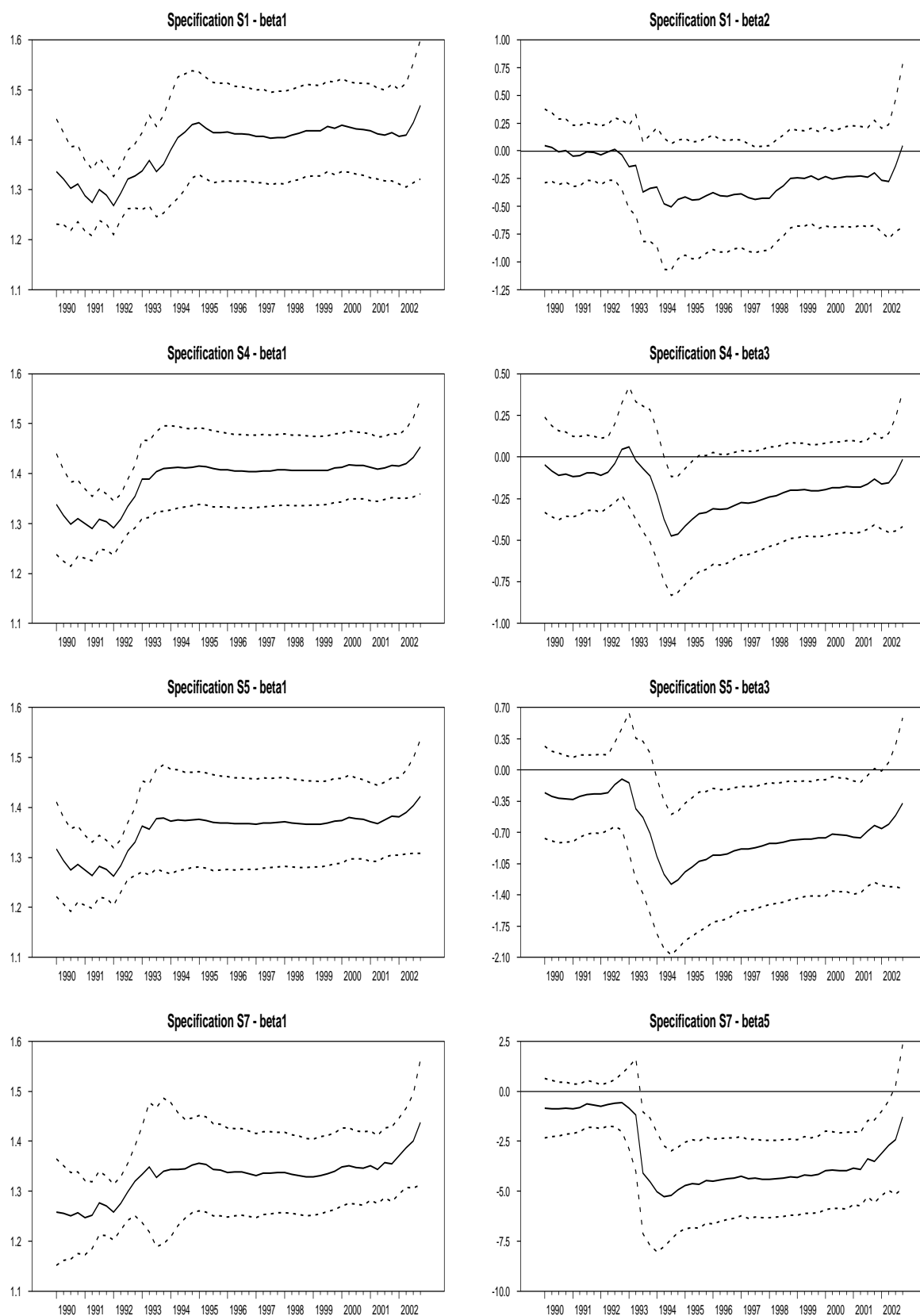


Fig. 1: Recursive FM-OLS parameter estimates for 1990Q1 to 2002Q4

main drawback of recursive parameter estimates. Moreover, the cointegration property might be lost after some point of time even with stable parameters simply because the disturbances of the long-run relationship are not stationary anymore.

Therefore, in a second step, we test for cointegration breakdown as outlined above using both OLS and FM-OLS to estimate the long-run relationship. We apply the test to the specifications S1, S4, S5 and S7 for the full sample from  $t = 1980Q1$  until  $t = T + m = 2002Q4$  with the length  $m$  of the post-breakdown sample varying from 1 to 16 quarters. The results of the  $P_c$  and the  $R_c$  tests for both the OLS and FM-OLS estimators are displayed in Figure 2 where the simulated  $p$ -values are graphed against the first period of the post-break sample  $T + 1$ . Note that the axis of  $p$ -values is rescaled such that the area of main interest, i.e.,  $p$ -values below 0.1 become better visible.

As a general result over all test statistics and specifications, stability is accepted at the start of the post-baseline period but is rejected at the end of this period. The  $p$ -values fall below the 10 percent line between 2000Q2 and 2001Q2, and below the 5 percent line between 2000Q4 and 2002Q1. Concentrating on specification S5 which yields the most plausible results for the baseline sample and is analyzed by Calza et al. (2001) and Bruggeman et al. (2003), we find that stability is rejected at the 5 percent level for all periods since 2001Q3 as indicated by the OLS  $P_c$  and  $R_c$  and by the FM-OLS  $R_c$  tests. This fits well with our hypothesis that excessive M3 growth, which started in the second half of 2001, may have caused a money demand instability. The test results for specifications S1 and S4 lead to the same conclusions while the test results for specification S7 suggests a slightly earlier instability.

A problem with these findings is that the cointegration-breakdown tests are designed for a structural change at a known breakpoint while so far we have rather used it as a tool to search for the break date. In a strict sense, this invalidates the critical values and the test overrejects the true null hypothesis of structural stability, as argued *inter alia* by Zivot and Andrews (1992). On the other hand, we constrained the time interval of a possible structural change to 1999Q1 to 2002Q4 which somewhat alleviates the overrejection problem. Moreover, if we simply perform the cointegration-breakdown tests at the break date of 2001Q4 which is a highly plausible candidate due to the start of excessive M3 growth,



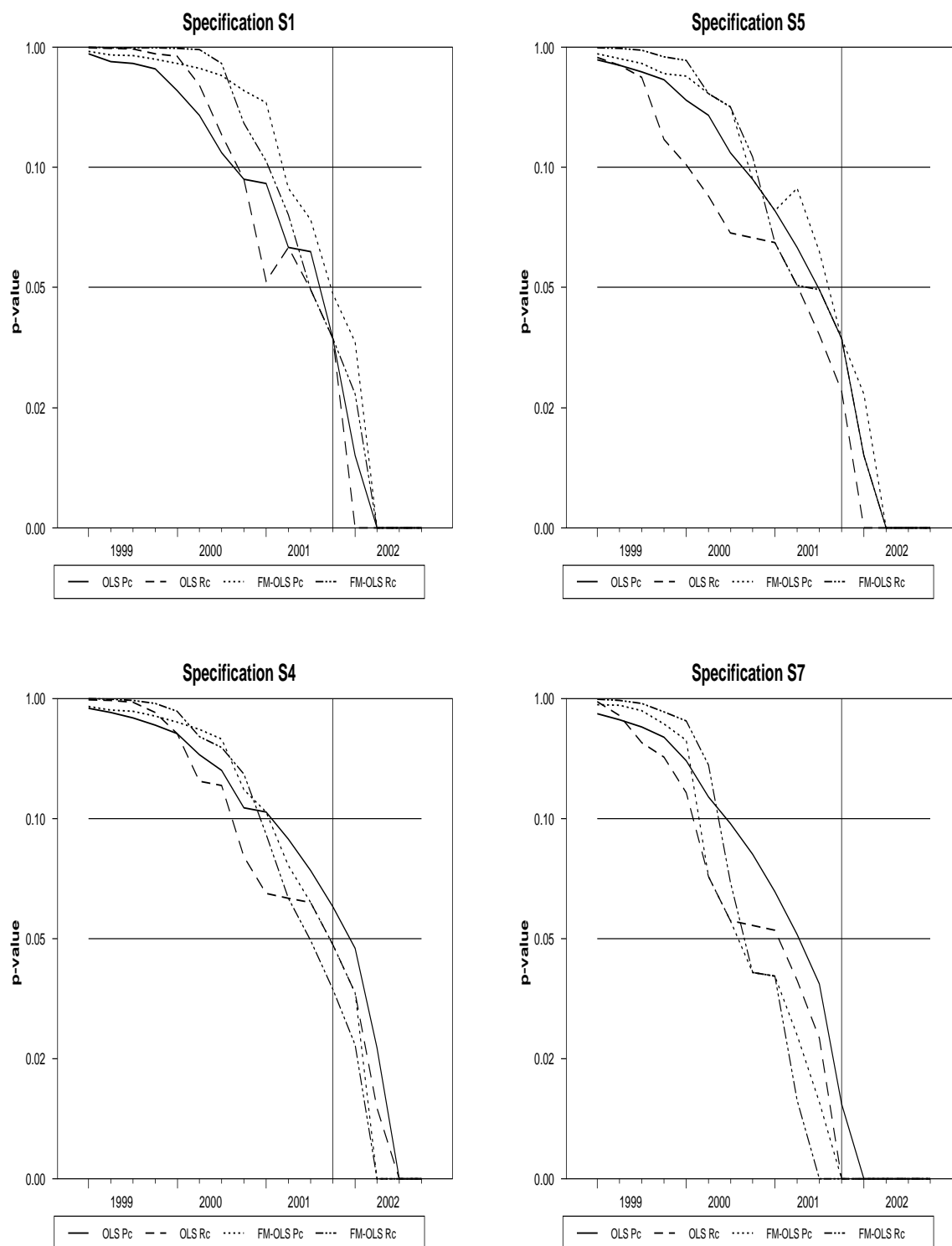


Fig. 2: Cointegration-breakdown tests for 1999Q1 to 2002Q4

we would again reject stability at the 5 percent level (see the vertical lines in Figure 2). It is therefore fair to conclude from the cointegration–breakdown tests that EMU money demand in fact has become unstable in the recent time.

In a next step, we ask whether the conventional Nyblom–type  $L_c$  test proposed by Hansen (1992b) for regressions with  $I(1)$  processes is able to detect the money demand instability. We choose the  $L_c$  test because most other stability tests suitable for cointegrating regressions require a trimming of the sample ends which precludes their use for the case at hand. The  $L_c$  test is designed to test the null hypothesis that the parameters are stable against the alternative that the parameters follow a martingale process and has the advantage that the time of the structural change is not pre–specified.

Applying the  $L_c$  test to all seven specifications yields the results presented in Table 3, first panel. Stability is only rejected for specification S6 contradicting the outcome of the cointegration–breakdown tests. There may be several reasons for this result. First, the specific alternative hypothesis of the  $L_c$  test implies that it is most appropriate for cases where the probability of structural changes is constant over the sample (Hansen, 1992b). However, in the case at hand we suspect an instability only at the end of the sample while the  $L_c$  test gives an average picture for the whole sample. Moreover, a simulation study by Andrews et al. (1996) indicates that power of the  $L_c$  test may be low for break points near the end of the sample. Finally, critical values derived from the asymptotic distribution of the test statistic may be a poor proxy for the small–sample critical values.

### 4.3 Modelling the long–run structural change

The test results presented in the previous section suggest that long–run structural stability of the EMU money demand function probably failed in the second half of 2001 and in 2002. It is therefore of interest to explain, or even model, this structural change. The ECB (2003a) argues that the increased uncertainty in equity markets has led to portfolio shifts from equities to safe and liquid assets which are part of M3. Bruggeman et al. (2003) use stock market volatility as explanatory variable in their analysis of EMU money demand. Since they assume that volatility is stationary, it can only be part of the short–run dynamics but not of the long–run relationship. As a consequence, it cannot be used

Tab. 3: Nyblom-type  $L_c$  tests for structural stability

<i>Tests for overall stability of the long-run (cointegration) relationships</i>							
Specification	S1	S2	S3	S4	S5	S6	S7
$L_c$	0.098	0.052	0.191	0.135	0.169	0.760**	0.123
<i>Tests for overall stability of the augmented long-run (cointegration) relationships</i>							
Specification	S1	S2	S3	S4	S5	S6	S7
$L_c$	0.484	0.310	0.572	0.249	0.265	0.956**	0.348
<i>Tests for overall stability of the short-run relationships (unrestricted ECM)</i>							
Specification	S1	S2	S3	S4	S5	S6	S7
$L_c$	5.91	6.25*	6.01*	5.96*	6.04*	6.00*	5.98*
<i>Tests for overall stability of the short-run relationships (parsimonious ECM)</i>							
Specification	S1	S2	S3	S4	S5	S6	S7
$L_c$	1.76	1.98	2.02	1.88	1.96	1.92	1.97
<i>Tests for individual parameter stability of the short-run relationships (parsimonious ECM)</i>							
Specification	S1	S2	S3	S4	S5	S6	S7
$L_c(\gamma_0)$	0.119	0.180	0.138	0.155	0.168	0.156	0.135
$L_c(\gamma_1)$	0.068	0.079	0.071	0.087	0.092	0.086	0.077
$L_c(\gamma_2)$	0.190	0.258	0.247	0.241	0.255	0.242	0.236
$L_c(\gamma_3)$	0.083	0.068	0.069	0.073	0.069	0.068	0.074
$L_c(\gamma_4)$	0.045	0.048	0.037	0.044	0.052	0.049	0.043
$L_c(\gamma_5)$	0.034	0.028	0.025	0.031	0.031	0.030	0.030
$L_c(\gamma_6)$	0.340	0.342	0.362*	0.332	0.328	0.334	0.349*
$L_c(\gamma_7)$	0.239	0.180	0.260	0.236	0.220	0.225	0.278
$L_c(\gamma_8)$	0.223	0.251	0.224	0.230	0.243	0.238	0.225
$L_c(\gamma_9)$	0.122	0.125	0.118	0.124	0.116	0.113	0.116
$L_c(\sigma^2)$	0.062	0.077	0.055	0.052	0.062	0.062	0.058

Notes: The critical values for the stability tests of the long-run relationships are 0.959 (1%), 0.623 (5%) and 0.497 (10%), cf. Hansen (1992b, Tab. 3). The critical values for stability tests of the augmented long-run relationships are 1.29 (1%), 0.901 (5%) and 0.752 (10%), cf. Hansen (1992b, Tab. 3). The critical values for the short-run stability tests are tabulated by Hansen (1990a) for up to 20 degrees of freedom. Since the unrestricted ECM contains 29 Parameters including the error variance, we simulated the asymptotic distribution given in Nyblom (1989, eq. 3.3) for this case. The critical values for the overall short-run stability tests (unrestricted ECM) are 7.00 (1%), 6.30 (5%) and 5.93 (10%). The critical values for the overall short-run stability tests (parsimonious ECM) are 3.27 (1%), 2.75 (5%) and 2.49 (10%). Finally, the critical values for the individual parameter stability tests are 0.748 (1%), 0.470 (5%) and 0.353 (10%). Test statistics being significant at the 10%, 5% or 1% level are denoted by \*, \*\* and \*\*\*, respectively.

to account for any long-run structural change in their setup.

Stationarity is, however, not an unquestionable property of stock market volatility. It is, e.g., a common finding in the empirical finance literature that volatility estimated from generalized autoregressive conditional heteroskedastic (GARCH) models is on the border to nonstationarity (Bera and Higgins, 1993, Bollerslev et al., 1992) and can, thus, be modelled as being integrated (Bollerslev and Engle, 1993, Engle and Bollerslev, 1986). We will leave the decision to several unit root tests which we apply to our measure of stock market volatility.

Additionally, money demand may also depend on real stock prices. For example, Choudhry (1996) finds evidence in favor of a long-run influence of real stock prices on Canadian and US demand for M1 and M2. For EMU money demand Kontolemis using data up to 2001Q3 finds a significant long-run influence of stock prices while Bruggeman et al. (2003) using data up to 2001Q4 obtain the opposite result. It is therefore unclear whether stock prices can account for the instability in the long-run money demand function. For our extended data set, we replicate their analysis. Because no European stock price index is available before 1987, we use the German DAX30 from 1980 to 1986 and the Dow Jones Euro Stoxx50 from 1987 to 2002. 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  $sp_t$ .

In a first step, we augment the seven specifications of money demand by our measure of real stock prices which yields specifications S1a to S7a. The estimation results are presented in the upper panel of Table 4. While the estimated income elasticities of money demand stay virtually unchanged, almost all other parameters are estimated very imprecisely. In particular, the influence of real stock prices on money demand is insignificant mimicking the results of Bruggeman et al. (2003).

Alternatively, we augment the seven specifications of money demand by a measure of stock market volatility  $v_t$ . To this end, we use the log of the quarterly median of squared daily yields which are calculated as log stock price changes. Applying the median instead of the mean function has the advantage that some huge outliers do not dominate the

Tab. 4: Estimates of the augmented money demand specifications

Specification	Estimated parameters						
	$\beta_1$	$\beta_2$	$\beta_3$	$\beta_4$	$\beta_5$	$\beta_6$	$\beta_7$
<i>With real stock prices</i>							
S1a	1.41 (0.091)	−0.38 (0.496)				−0.002 (0.0279)	
S2a	1.47 (0.123)	−0.04 (0.855)		0.04 (0.855)		−0.000 (0.0272)	
S3a	1.40 (0.121)	−0.48 (0.863)		0.48 (0.863)	−0.99 (2.148)	0.015 (0.0268)	
S4a	1.45 (0.074)		−0.33 (0.283)			−0.015 (0.0221)	
S5a	1.42 (0.094)		−0.79 (0.608)	0.79 (0.608)		−0.010 (0.0236)	
S6a	1.41 (0.077)		−1.79 (0.613)	1.79 (0.613)	2.99 (1.761)	−0.015 (0.0190)	
S7a	1.38 (0.114)				−1.15 (2.135)	0.014 (0.0275)	
<i>With stock market volatility</i>							
S1b	1.41 (0.068)	−0.24 (0.357)					−0.003 (0.006)
S2b	1.49 (0.072)	0.71 (0.713)		−0.71 (0.713)			0.005 (0.007)
S3b	1.42 (0.065)	0.58 (0.609)		−0.58 (0.609)	−3.13 (1.495)		−0.001 (0.006)
S4b	1.42 (0.046)		−0.13 (0.213)				0.001 (0.005)
S5b	1.41 (0.056)		−0.37 (0.484)	0.37 (0.484)			0.002 (0.006)
S6b	1.40 (0.052)		−0.24 (0.591)	0.24 (0.591)	−0.56 (1.733)		0.004 (0.005)
S7b	1.39 (0.054)				−2.53 (1.479)		−0.001 (0.006)
<i>With real stock prices and stock market volatility</i>							
S1c	1.46 (0.067)	−0.38 (0.353)				−0.023 (0.020)	0.009 (0.0052)
S2c	1.54 (0.083)	0.25 (0.562)		−0.25 (0.562)		−0.029 (0.020)	0.016 (0.0059)
S3c	1.47 (0.068)	0.61 (0.476)		−0.61 (0.476)	−2.52 (1.215)	−0.013 (0.017)	0.009 (0.0049)
S4c	1.48 (0.056)		−0.37 (0.205)			−0.032 (0.017)	0.009 (0.0045)
S5c	1.47 (0.069)		−0.70 (0.431)	0.70 (0.431)		−0.031 (0.018)	0.012 (0.0053)
S6c	1.45 (0.065)		−0.71 (0.528)	0.71 (0.528)	0.10 (1.550)	−0.025 (0.017)	0.010 (0.0053)
S7c	1.44 (0.066)				−1.99 (1.243)	−0.017 (0.018)	0.009 (0.0053)

Notes: All specifications are estimated by FM-OLS with automatic bandwidth selection. The parameter  $\beta_6$  measures the influence of real stock prices  $sp_t$ ,  $\beta_7$  measures the influence of stock market volatility  $v_t$ . Standard errors are reported in brackets below the estimates.

Tab. 5: Unit root tests for stock market volatility

Test	Lags for autocorrelation correction				critical values		
	1	2	3	4	1%	5%	10%
ADF	-2.40	-1.92	-1.69	-1.69	-3.51	-2.89	-2.58
DFGLSu	-2.48	-2.04	-1.85	-1.87	-3.28	-2.73	-2.46
KPSS	1.25	0.94	0.77	0.66	0.74	0.46	0.35

Notes: ADF is the augmented Dickey–Fuller test, DFGLSu is the Dickey–Fuller test with GLS detrending proposed by Elliott (1999), KPSS is the stationarity test proposed by Kwiatkowski et al. (1992). All test regressions are estimated with a constant but without a trend.

volatility measure of a whole quarter. Before using volatility as an explanatory variable for money demand, we employ three different unit root tests to determine whether it is stationary or not. Both the augmented Dickey–Fuller (ADF) test and the DFGLSu test (Elliott, 1999) have the null hypothesis of a unit root while the KPSS test has the null hypothesis of stationarity. Autocorrelation tests indicate that one lagged difference in the ADF test equation is sufficient to guarantee uncorrelated errors. The test statistic of  $-2.40$  is not even significant at the 10% level. Thus, nonstationarity cannot be rejected. This result is not sensitive to adding more lags. Also, using the DFGLSu test instead which employs a GLS detrending and is more efficient than the ADF test, does not change the result. Applying the KPSS test, stationarity has to be rejected at the 5% level for all lags. As a consequence, we assume in the following that stock market volatility behaves like a nonstationary variable and may therefore well be part of a cointegrating relationship.

Estimating the money demand specifications augmented by stock market volatility, S1b to S7b, yields the results given in the middle panel of Table 4. Unfortunately, the influence of volatility turns out to be insignificant. The same result is obtained by Bruggeman et al. (2003) who do not find any explanatory power of stock market volatility as a stationary variable in a cointegrated VAR model of EMU money demand. They even report that the inclusion of the volatility variable induces instability instead of absorbing it.

Finally, we add both real stock prices and stock market volatility giving rise to money demand specifications S1c to S7c. This may be necessary because the influence of stock markets on money demand may be only measurable if both stock prices and volatility are high so that a huge amount of funds flows into safe and liquid instruments which are part

of M3. The estimation results are presented in the lower panel of Table 4. Especially our preferred specifications S1c, S4c, S5c and S7c now exhibit plausible and significant parameter estimates. Both a decline in real stock prices and a rise of stock market volatility leads to an increase in money demand.

Using a money demand functions augmented with real stock prices and stock market volatility is only a viable alternative to standard specifications if it solves the instability problem. To this end, we again carry out cointegration–breakdown and  $L_c$  tests. The  $p$ -values of the cointegration–breakdown tests are presented in Figure 3. At the 5% level, stability is now only rejected for the last observation and even this rejection is not clear-cut. In specification S4, stability is never rejected while in specifications S1 and S5 stability is not rejected at the 5% level when the testing procedure is applied to FM–OLS estimates. In comparison to the test results for the baseline specifications in Figure 2, the instability problem is at least greatly reduced. The  $L_c$  test is reported in the second panel of Table 3. Instability is only detected for specification S6 but not for our preferred specifications. We therefore conclude that there is not enough evidence that the augmented money demand functions are unstable even if the results are not clear-cut for the last observation.

#### 4.4 Tests for short–run stability

In a next step, we check whether the money demand specifications also exhibit short–run stability. To this end, we set up an error–correction model (ECM) for each specification taking the estimated long–run parameters as given. This is asymptotically justified by their superconsistency. Therefore, we proceed as follows for each specification. We estimate the long–run parameters by FM–OLS, calculate the cointegration residuals  $\hat{u}_t^s$  for each specification  $s = \text{S1c}, \dots, \text{S7c}$ , and plug them into an ECM with  $\Delta mp_t$  as dependent variable and the following explanatory variables:  $\Delta mp_{t-1}$ ,  $\Delta y_t$ ,  $\Delta r_t^l$ ,  $\Delta r_t^s$ ,  $\Delta r_t^o$ ,  $\Delta^2 p_t$ ,  $\Delta sp_t$ ,  $\Delta v_t$  and  $\Delta oil_t$  which is a world oil price index taken from Datastream.

To choose an appropriate lag order, we perform LM–tests for autocorrelation which indicate that a minimum of two lags is necessary to guarantee uncorrelated disturbances.

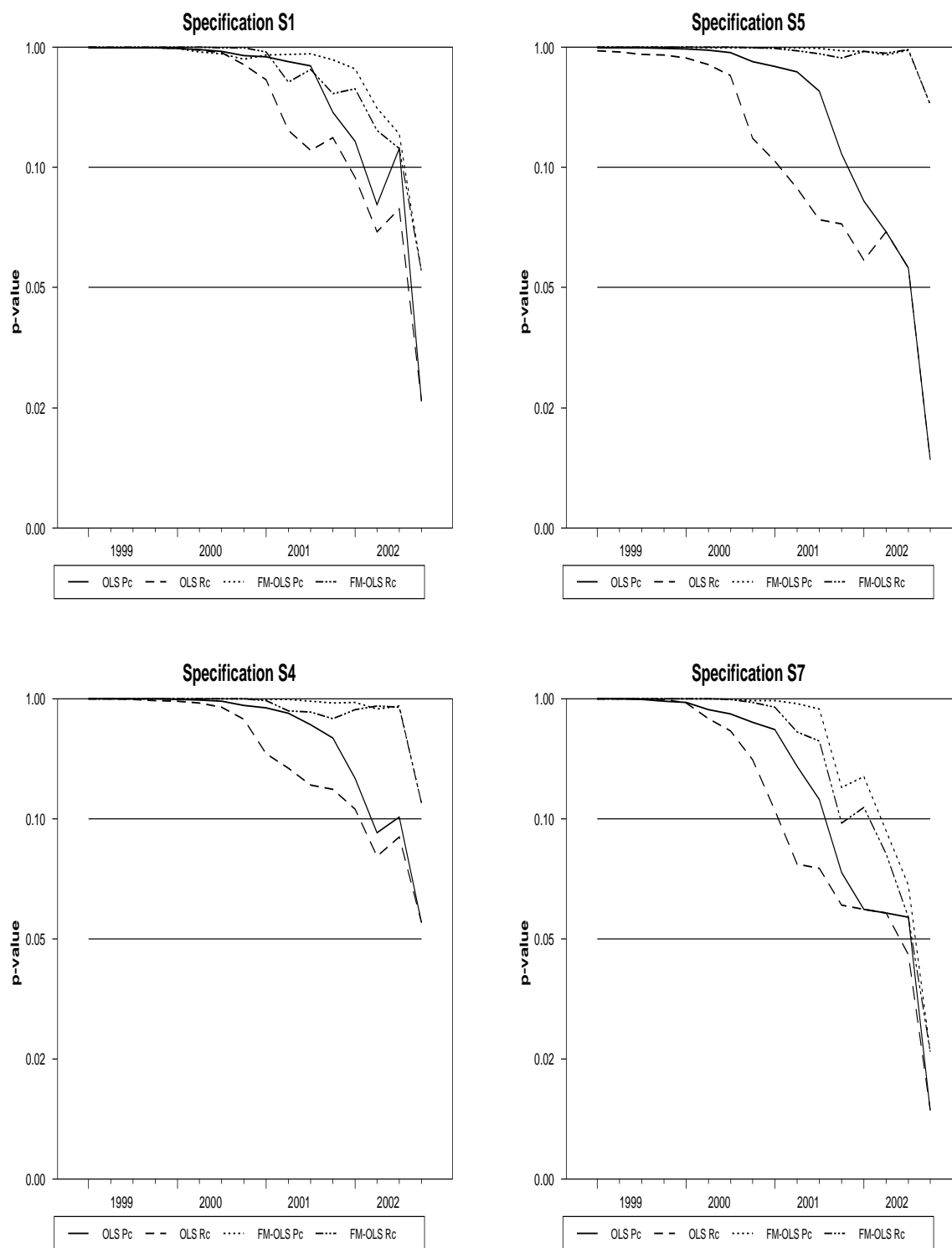


Fig. 3: Cointegration-breakdown tests for 1999Q1 to 2002Q4



As a result, the unrestricted ECM is given as

$$\begin{aligned}
\Delta mp_t = & \alpha_0 + \alpha_1 \hat{u}_{t-1}^s + \alpha_2 \Delta mp_{t-1} + \alpha_3 \Delta mp_{t-2} + \alpha_4 \Delta y_t + \alpha_5 \Delta y_{t-1} + \alpha_6 \Delta y_{t-2} + \alpha_7 \Delta r_t^l \\
& + \alpha_8 \Delta r_{t-1}^l + \alpha_9 \Delta r_{t-2}^l + \alpha_{10} \Delta r_t^s + \alpha_{11} \Delta r_{t-1}^s + \alpha_{12} \Delta r_{t-2}^s + \alpha_{13} \Delta r_t^o + \alpha_{14} \Delta r_{t-1}^o \\
& + \alpha_{15} \Delta r_{t-2}^o + \alpha_{16} \Delta^2 p_t + \alpha_{17} \Delta^2 p_{t-1} + \alpha_{18} \Delta^2 p_{t-2} + \alpha_{19} \Delta sp_t + \alpha_{20} \Delta sp_{t-1} \\
& + \alpha_{21} \Delta sp_{t-2} + \alpha_{22} \Delta v_t + \alpha_{23} \Delta v_{t-1} + \alpha_{24} \Delta v_{t-2} + \alpha_{25} \Delta oil_t + \alpha_{26} \Delta oil_{t-1} \\
& + \alpha_{27} \Delta oil_{t-2} + \varepsilon_t
\end{aligned} \tag{8}$$

with 28 free parameters. Since many of the estimated parameters turn out to be insignificant, the subsequent stability tests may lack power. We therefore also estimate a parsimonious ECM from which all insignificant parameters are deleted in a stepwise procedure. We end up at an identical lag structure for all specifications,

$$\begin{aligned}
\Delta mp_t = & \gamma_0 + \gamma_1 \hat{u}_{t-1}^s + \gamma_2 \Delta mp_{t-1} + \gamma_3 \Delta r_{t-1}^l + \gamma_4 \Delta r_{t-1}^o + \gamma_5 \Delta^2 p_t + \gamma_6 \Delta^2 p_{t-1} \\
& + \gamma_7 \Delta sp_{t-1} + \gamma_8 \Delta v_t + \gamma_9 \Delta oil_t + \varepsilon_t,
\end{aligned} \tag{9}$$

where only 10 parameters are left.

The  $p$ -values of the end-of-sample stability test for stationary regressions applied to our preferred specifications S1, S4, S5 and S7 are presented in Figure 4. Obviously, the estimated  $p$ -values are similar for all four specifications and far above any conventional significance level. Consequently, the null hypothesis of structural stability cannot be rejected.

Like before, we supplement the end-of-sample stability tests by a series of Nyblom-type  $L_c$  tests for structural stability which are proposed by Hansen (1992a) for stationary regressions. These tests are both applied to each individual parameter and to the whole ECM, again taking the cointegration parameters as given. Test statistics and asymptotic critical values are presented in Table 3. For the unrestricted ECM the overall stability tests are significant at the 10% but not at the 5% level. Together with the results of the end-of-sample stability tests, this is not enough evidence against instability. The results for the parsimonious ECM are even more favorable. The overall stability test are insignificant at

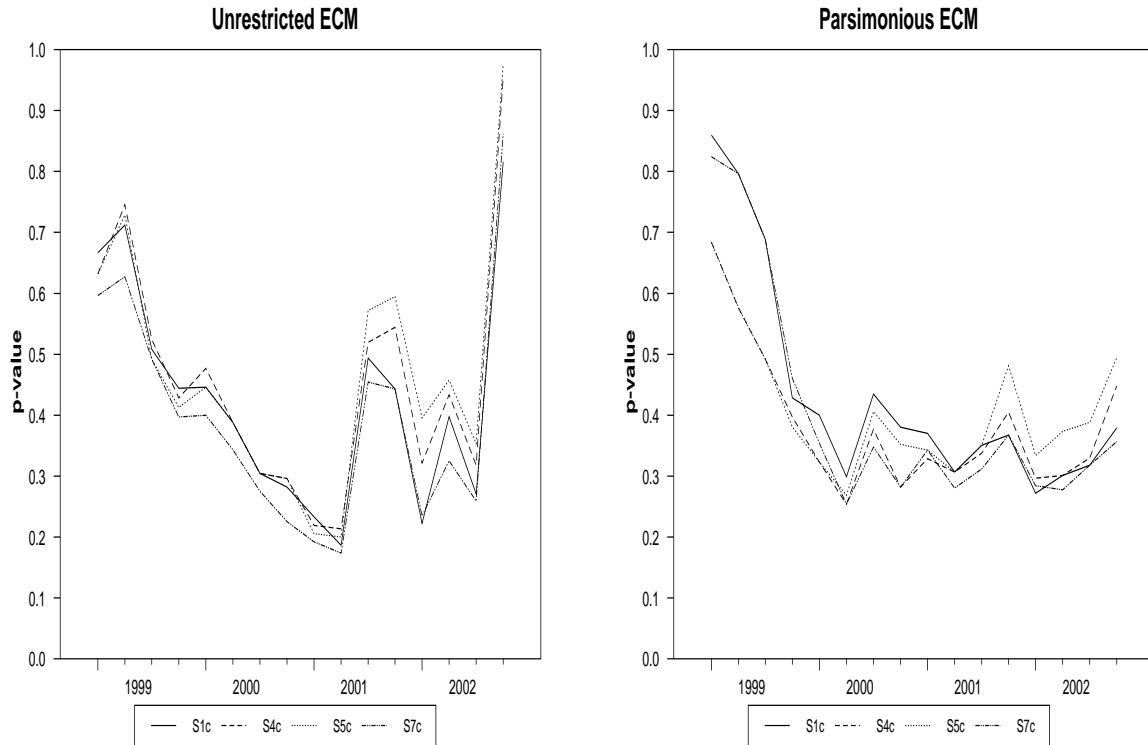


Fig. 4: Short-run stability tests for 1999Q1 to 2002Q4

the 10% level. Consequently, stability cannot be rejected. Even if looking at each single parameter, stability cannot be rejected at the 5% level.

## 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 an estimated money demand function, we are now in a position to quantify the portion of money demand which is due to stock market developments. To this end, we take all variables except for nominal M3 as exogenous and forecast M3 series conditional on a hypothetical path of real stock prices and stock market volatility and on the actual paths of all other variables. Since the model is dynamic, the forecasted series are calculated from dynamic simulations and are (slightly) different from the fitted series. To assess the portion of M3 growth which is due to stock market developments, we therefore perform a reference simulation which is conditional on the actual paths of all variables

including stock prices and volatility. The difference between these two simulation is then exclusively related to the different assumptions regarding the stock market developments.

We analyze three different scenarios of stock market developments. In the first scenario, we assume that both stock market prices and volatility remained at their 2000Q3 levels when stock market prices were on an all-time high and uncertainty was still low. From this scenario we can learn how money demand would have developed if no stock market downswing had occurred. In the second scenario, stock market conditions are those of the first quarter 1999 when EMU started. This scenario helps us infer the development of money demand if stock market conditions had stayed unchanged during the time the ECB has been responsible for monetary policy. Finally, in a third scenario stock market prices and volatility from 1998Q1 onwards are set to their average values of the period 1990:Q1 to 1997Q4. This scenario tells us how money demand would have developed if neither the rise nor the burst of the stock market bubble had occurred. Of course, one has to bear in mind that we do not change the paths of the remaining exogenous variables GDP, interest rates, inflation and oil prices. The experiments therefore neglect any interrelations between these variables and the stock market developments. 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 answer the question how the stock market developments *ceteris paribus* have affected money demand.

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 Figure 5 together with actual M3 growth rates. The first scenario is given in the upper left panel. Actual and simulated M3 growth rates are by assumption identical up to 2000Q3. Afterwards, scenario 1 simulation assumes constant stock market conditions. As a consequence, money demand remains low and M3 growth rates continue to fluctuate around 4%. In contrast, the reference simulation, which uses the actual stock market developments as input, leads to rising money demand and M3 growth rates around 6.0% since 2001Q4. The difference between these two simulations is displayed in the lower right panel. It suggests that since 2001Q4 roughly 2 percentage points of M3 growth have been due to the stock market downswing and increased volatility. This implies that actual M3

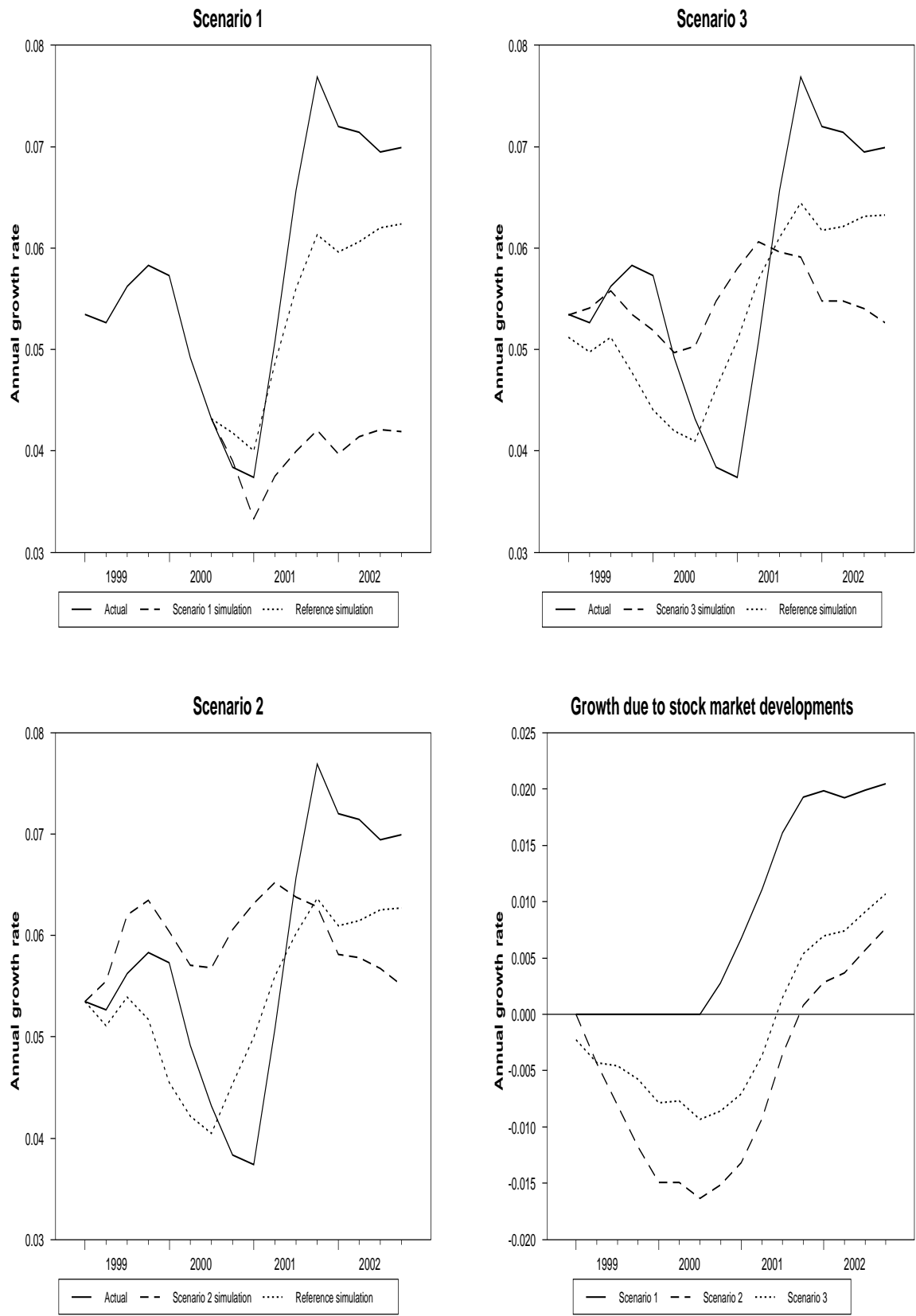


Fig. 5: Simulated annual money growth for 1999Q1 to 2002Q4

growth rates which have been between 7% and 8% in this period, would have been reduced to the range between 5% to 6% without the adverse stock market developments. Thus, the reference growth rate of 4.5% announced by the ECB would have still been exceeded but by far less than actually observed. Given that the funds shifted from stock markets to M3 are held for speculative purposes only and, hence, will be shifted back towards stock markets once the conditions develop more favorably again, potential inflationary pressure arises only due to excess M3 growth of 0.5 to 1.5 percentage points in contrast to 2.5 to 3.5 percentage points without corrections for stock market developments. From this perspective, we may thus conclude that the ECB has good reason to argue that the actually observed high M3 growth rates do not pose a major threat to price stability and that interest rates may remain at their low level.

However, a strong caveat applies. The conclusions drawn from the first simulation rest upon the implicit assumption that stock market prices will soon return to their 2000Q3 level. At the moment, this is rather unlikely to happen. On the contrary, one might well argue that the 2000Q3 level resulted from stock market exaggerations only and led to particularly low money demand which was unforeseeable at the start of EMU. From this perspective, unexpectedly favorable stock market conditions may have supported acceptable M3 growth rates in the first two years of EMU. To analyze this, we perform a second simulation of money demand keeping real stock prices and volatility at their 1999Q1 levels, see Figure 5, lower left panel. A comparison to the reference simulation which uses actual stock prices and volatility as input shows that the stock market boom has in fact caused a reduction of M3 growth in 1999 and 2000. The difference between scenario 2 simulation and the reference simulation is less than  $-1.5$  percentage points in 2000Q3. This implies that without the stock market boom M3 growth would have been considerably higher than actually observed. This is reversed only in 2002 when the difference between scenario 2 simulation and the reference simulation rises up to 0.75 percentage points. We can therefore conclude that compared to the conditions at the start of EMU, stock market developments can only be blamed for only a small portion of the actually high M3 growth rates.

One could even argue that the whole stock market boom is an exceptional event which should not be taken into consideration when calculating “normal” M3 growth. In a third ex-

periment we therefore simulate M3 growth given the stock market conditions since 1998Q1 equal the average conditions between 1990Q1 and 1997Q4.<sup>3</sup> They differ from the stock market conditions in 1999Q1 when both real stock prices and volatility were higher than in the beginning of the nineties. The simulation results which are displayed in Figure 5, upper right panel, are qualitatively similar to simulation 2. In the first two years of EMU, favorable stock market conditions helped keeping M3 growth rates rather low while adverse stock market conditions have driven M3 growth rates up since the end of 2001. Quantitatively, the stock market development accounts for up to slightly more than 1 percentage point of the excessive M3 growth rates in 2002 which is again a rather small portion.

Simulations 2 and 3 indicate that compared to non-exaggerated or average stock market conditions, the stock prices and volatility in 2002 account only to a small extent for the high M3 growth rates. This in turn implies that at least a non-negligible portion of the growth rates exceeding the reference value of 4.5% may translate into prices changes. It is thus not at all obvious that the M3 growth rates do not pose any threat to price stability if the economy remains at its current level.

## 6 Conclusion

This paper has analyzed the stability properties of various money demand specifications proposed in the literature. They have all 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 had to be rejected for these specifications. The tests indicated that the break point is probably in the year 2001 when M3 growth increased and stock market conditions deteriorated.

In an effort to restore a stable relationship between money and prices, we augmented the conventional money demand functions with two variables from the financial sector: real stock prices and stock market volatility. It turned out that these augmented specifications exhibit much better stability properties than the conventional ones. In particular, the tests based on FM-OLS estimates of the cointegration relationship did not reject long-run

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<sup>3</sup> This implies the assumption that the stock market boom started in 1998Q1 which is certainly difficult to defend. However, changing the starting point by some quarters does not change the overall conclusions.

stability anymore. Moreover, short-run stability was tested, and could not be rejected, with the help of the end-of-sample stability test introduced by Andrews (2002) applied to single-equation error-correction models which were estimated conditional on the long-run parameters.

Given stability holds for the augmented specifications we analyzed the impact of the stock market developments on M3 growth. To this end, we used a stable money demand function to perform dynamic simulations of M3 growth conditional on different hypothetical paths of stock prices and volatility. The simulations suggest that, compared to the all-time high in 2000Q3, the stock market conditions in 2002 accounted for roughly 2 percentage points of the M3 growth rates. This seems to support the position of the ECB that a major portion of the excessive M3 growth rates is due to adverse financial developments and, hence, will not pose a threat to price stability.

However, this argument depends on choosing 2000Q3 as the reference point to which stock markets will return in the near future. If instead the stock market boom is interpreted as an exceptional event and the reference point is chosen as the first period of EMU, the conclusions are quite different: In the first two years of EMU the rise in stock prices led to unusually low M3 growth rates which spuriously signalled that price stability is not threatened. While stock market conditions fell below the reference values at the end of 2001, a much smaller portion of the 2002 M3 growth rates can be traced back to adverse financial developments. From that point of view, the ECB should at least be cautious when it argues that the high M3 growth rates will not transmit into inflation in the medium run.

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