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Investigating M3 money demand in the euro area. New evidence based on standard models

Christian Dreger and Jürgen Wolters[1]

Abstract: Monetary growth in the euro area has exceeded its target level especially since 2001. Likewise, recent empirical studies did not find evidence in favour of a stable long run relationship between the variables entering the money demand function. Instead the equation appears to be increasingly unstable if more recent data are included. Since the link between money balances and macroeconomic variables seems to has become rather fragile, these results put serious doubts concerning the rationale of monetary aggregates in the monetary policy strategy of the ECB. However, if the analysis is done without imposing a short run homogeneity restriction between money and prices, a stable long run money demand relationship can be identified, where recursively estimated parameters are almost stable. In addition, the corresponding error correction model survives a wide array of specification tests, including procedures for nonlinearities and parameter instability. Hence, the apparent monetary overhang is in line with standard models of money demand behaviour, and is not expected to lead to a rise in inflation.

Keywords: Cointegration analysis, error correction, money demand, monetary policy

JEL Classification: C22, C52, E41

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1 Introduction

Monetary aggregates play a crucial role in the alignment of the two pillar monetary policy strategy of the ECB. While one pillar is based on the economic analysis of price risks in the short term, the other pillar includes the monetary analysis of risks to price stability especially in the medium and long term (ECB, 2003). The reference value for monetary growth is taken as a benchmark for assessing monetary developments. Since the end of 2001, M3 reference growth rates have been continuing to exceed the target of about 4.5 percent by more than 2.5 percentage points. But, inflation did not accelerate at all, thereby questioning the rationale of monetary aggregates in the monetary policy strategy of the ECB. If the link between money and prices is increasingly unstable, money growth is not a well-designed tool to analyze future inflation prospects and support policy decisions.

For monitoring the inflation process, a stable money demand function is extremely important, at least as a long run reference. If this condition is met, money demand can be linked to the real side of the economy. Variables explaining money demand include income, interest and inflation rates. However, recent evidence has raised serious doubts concerning the robustness of the relationship. If data up to 2001 are used, conventional money demand functions for the euro area can be firmly established, see Fagan and Henry (1998), Funke (2001), Coenen and Vega (2001), Bruggemann, Donati and Warne (2003), Brand and Cassola (2004) and Holtemöller (2004a, b). Extending the sample to a more recent period usually destroys these findings, as cointegration between the variables cannot be detected any longer, see Gerlach and Svensson (2003), Carstensen (2004) and Greiber and Lemke (2005). This has led some authors to focus on the relationships not between the original variables, but between the core components, either generated by HP filtered values or moving averages, see Gerlach (2004) and Neumann and Greiber (2004). In other studies, measures of uncertainty are allowed to enter the long run equation. Using the extended model, Carstensen (2004) and Greiber and Lemke (2005) are able to find support for a money demand relationship. However, in principle, proxies for uncertainty should be stationary, implying that this approach is not really convincing. Brüggemann and Lütkepohl (2005) have found stable money demand relationships using data up to 2002. In contrast to all other papers on euro area money demand, they employed German data until the end of 1998.

Despite the results from the previous literature, this paper presents evidence in favour of a stable long run money demand relationship specified in the observables. The inclusion of the inflation rate is crucial for the existence of the long run equation. Inflation might capture opportunity costs of holding real assets and can be also related to portfolio adjustment processes. In addition, two impulse dummies are introduced. While the first one (1990.2) refers to the German unification, the other one (2001.1) points to the burst of the stock market bubble. The money demand relationship is stable, as shown by recursive estimates of the cointegration space, and the corresponding error correction model resists a battery of specification tests. These include a number of recent tests for parameter instability and nonlinearities based on the smooth transition regression approach (Granger and Teräsvirta, 1993 and Terävista, 2004).

The rest of the paper is organized as follows. Section 2 reviews the specification of the long-run money demand function. In section 3, the data series used in the empirical analysis are discussed. Specification and estimation of money demand functions in error correction form has been the customary approach to capture the nonstationary behaviour of the time series involved. Robust evidence regarding the cointegration properties is provided in section 4. In section 5, an error correction model for money demand is estimated and tested for a broad range of misspecification. Section 6 concludes.

2 Specification of money demand

In this paper, a widely used specification of money demand is chosen as the point of departure. According to Ericsson (1998), the bulk of theories of money demand behaviour imply a long run relationship of the form

$$(1) \qquad M / P = f(Y, OC)$$

where M denotes nominal money, P price level, Y income, representing the transaction volume in the economy and OC a vector of opportunity costs of holding money. Price homogeneity is assumed to be valid as a long-run condition. In fact, the money stock and the price level might be I(2) variables. If these variables are cointegrated, real money balances could be I(1). In this case the long run homogeneity restriction maps the money demand analysis into an I(1) system, see Holtemöller (2004b). According to textbook presentations, the scale variable is expected to exert a positive effect on nominal and real money balances. Typical models in the literature differ in the concrete specification of opportunity costs, see Golinelli and Pastorello (2002) for a survey. If the costs measure the earnings of alternative financial assets, possibly relative to the own yield of money balances, their coefficients should enter with a negative sign. Inflation is usually interpreted as a part of the opportunity costs, as it represents the costs of holding money in spite of holding real assets, see Ericsson (1998). But, the inclusion of inflation can also be justified by different arguments. In presence of adjustment costs and nominal inertia, Wolters and Lütkepohl (1997) have shown that the variable should enter the long run relationship for real balances, even if it does not enter the equation for nominal balances. Hence, inflation allows to discriminate whether the adjustment process is in nominal or real terms (Hwang, 1985). Alternatively, the inclusion of the inflation rate provides a convenient way to generalize the short run homogeneity restriction imposed between money and prices. While the restriction is justified from a theoretical point of view, there might be a lack of support in the particular observation period.

Usually, a semi logarithmic linear specification of long run money demand is preferred in the empirical analysis

$$(2) \qquad m_t - p_t = \delta_0 + \delta_1 y_t + \delta_2 r_t + \delta_3 \pi_t$$

where *m-p* is log real money balances, *y* is log of real income, *r* the nominal return of financial assests and π the annualized inflation rate, i.e. $\pi = 4\Delta p$ if quarterly data are used, and *t* the time index. The parameters $\delta_1 > 0$, $\delta_2 < 0$, and δ_3 denote the income elasticity, and the semielasticities with respect to the return of other financial assets and inflation, respectively. Due to the ambuigity in the interpretation of the inflation variable, the sign of its impact cannot be specified on theoretical reasoning.

3 Data and preliminary analysis

With the introduction of the euro on January 1, 1999 the responsibility for monetary policy was transferred to the ECB. As the time series under the new institutional framework are too short to draw any robust conclusions, they have to be extented by artificial data. Usually, euro area series prior to 1999 are obtained by aggregating national time series, see for example Artis and Beyer (2004). But, different aggregation methods are available and can lead to different results. By comparing aggregation approaches based on methods using variable or fixed period exchange rates, Bosker (2006) has emphasized that the differences between both methods for money demand variables are substantial prior to 1983, in particular for interest and inflation rates. However, they are almost negligible from 1983 onwards. In addition, the European Monetary System started working in 1983, and the financial markets of the member countries were much more integrated than before. Therefore, the observation period in this study is 1983.1-2004.4, where quarterly seasonally adjusted series are employed.

Nominal money balances are taken from the ECB monthly bulletin database and refer to M3 and end of period values. The short and long term interest rates *rs* and *rl* are also obtained from this source and defined by the end of period 3month Euribor and 10 years government bond rate, respectively. Income is proxied by nominal GDP taken from Eurostat as well as the GDP deflator (1995=100). Both series begin in 1991.1. Due to evidence presented by Holtemöller (2004a), the Brand and Cassola (2004) GDP data should be used in the earlier period, as these data yield stable and economically interpretable results. Note that this choice does not affect any conclusions in this paper, as the instability of money demand is only a problem in recent years. All variables are in logarithms with the exception of interest rates. Inflation is defined as $\pi=4\Delta p$, where p is the logarithm of the GDP deflator and Δ the first difference operator. In order to obtain real money balances (m-p) and real income (y), the nominal series are deflated with the GDP deflator (1995=100). Figure 1 displays the evolution of series in levels in the 1983.1-2004.4 period, while figure 2 shows the first differences.

-Figures 1 and 2 about here-

Several comments are in order. First, the variables considered are integrated of order 1, I(1), that is, they are nonstationary in the levels representation, but stationary in first differences. This well-known result has been reported in numerous empirical studies, see Coenen and Vega (2001), Golinelli and Pastorello (2002) and Holtemöller (2004a, b), among others. The results of the integration tests are omitted here to save space, but can be obtained from the authors upon request. Second, outliers occur in real money balances, see the graph for the first differences. The first one (1990.2) is due to the German unification, while the other one (2001.1) refers to the eve of stock market turbulences, see Kontolemis (2002). In particular, the large decrease in stock markets led to a higher demand for liquid assets. In the subsequent analysis, these outliers are acknowledged by means of two impulse dummies, which are equal to 1 in the respective period and 0 otherwise (d902 and d011).

4 Cointegration analysis

In systems including real money balances, real income, the interest rate and inflation, at least one cointegration relationship should represent a long run money demand equation in the style of (2). Furthermore, there is a second possible cointegration vector. Examining German data, Hubrich (2001) and Lütkepohl and Wolters (2003) have detected a stationary real interest rate due to the Fisher effect, i.e. a relation between the nominal interest rate and inflation. To investigate the cointegration properties for different systems of variables, the Johansen (1995) trace test is employed as the workhorse, see table 1 for the results. Estimation is done using Eviews 5.1. To correct for finite samples, the trace statistic is multiplied by the scale factor $(T\text{-}pk)/T$, where T is the number of the observations, p the number of the variables and k the lag order of the underlying VAR model in levels (Reimers, 1992). The lag length of the VARs is estimated using the Schwarz criterion. The constant enters in an unrestricted way, together with the two impulse dummies for 1990.2 and 2001.1 needed for variable sets including real money balances. However, if linear trends can be safely excluded, the constant is restricted to the cointegration vector, and dummies are not involved. This is the case for systems comprising only interest rates and inflation.

-Table 1 about here-

There is a strong indication for exactly one cointegrating relation in the (*m-p*, *y*, *rs*, *π*), (*m-p*, *y*, *π*) and (*rs*, *π*) system, respectively. This evidence could be consistent with a money demand relationship in the long run, possibly excluding the interest rate, and the Fisher equation. However, a second cointegration relation in the four variable system is supported only at the 0.2 level of significance.

Regarding money demand, the two sets (*m-p*, *y*, *rs*, Δ*p*) and (*m-p*, *y*, Δ*p*) are of interest. The exclusion of the interest rate from the former is supported by a likelihood ratio test (chi square 1.37, *p*-value 0.24). The implied cointegration relationships can be normalized on real money balances,

(3a) $(m\text{-}p, y, \pi)$: $ec1=(m\text{-}p)-1.238y+5.162\pi$

(3b) $(m\text{-}p, y, rs, \pi)$: $ec2=(m\text{-}p)-1.266y+4.528\pi$

and are almost perfectly correlated over the observation period. For reasons of a parsimonous model, the error correction term (3a) is favoured as the base for the subsequent analysis. However, all results remain valid with the alternative (3b). It is very remarkable that similar cointegrating relationships have been found by Wolters, Teräsvirta and Lütkepohl (1998) and Lütkepohl and Wolters (2003) for the German economy. The mean-adjusted deviations from the long run relation are displayed in figure 3. No abnormal behaviour can be detected in the series even after 2001. To complete the discussion, the cointegration evidence in the (*rs*, *π*) system is only a weak support for the Fisher equation. In fact, the null that the cointegration parameters of inflation and interest rates are equal but of opposite sign, leads to a chi-square test statistic of 3.66, with a *p*-value of 0.06.

-Figure 3 and table 2 about here-

To gain insights into the stability of the cointegration property and the long run vector in the (*m-p*, *y*, *π*) system, recursive estimation techniques are applied. Table 2 exhibits the results from this exercise, where the trace statistic and the cointegration vector are estimated using forward and backward methods. Furthermore, the cointegration vector is estimated by means of a moving window approach. Overall, the relationships seem to be very stable, even in the pretented instability period after 2001. In particular, the cointegration finding can be confirmed in any case. There might be an upward shift in the inflation coefficient in absolute value towards the end of the sample, but this change is hardly significant.

5 Error correction modeling

Whether or not the cointegrating relationship can be interpreted in terms of a money demand function is inferred from the error correction model. However, as we are mostly interested in the stability of a money demand equation, the analysis is concentrated on conditional single equation models. Here, the cointegration vector according to the specification (3a) is employed. At the initial stage of the estimation process, in addition to the error correction term the contemporaneous and the first two lags of the changes of all variables, a constant and the two impulse dummies are included. In addition, it is also tested whether the short run dynamics are influenced by wealth effects, arising from the stock and housing markets (e.g. Kontolemis, 2002). In the first round, those variables could control for possible short run instabilities in money demand. Then, the variables with the lowest and insignificant *t*-values have been eliminated successively (0.1 level). The final money demand relationship is (*t*-values in parantheses)

$$\text{(4)} \quad \begin{aligned} \Delta(m-p)_t = & -\underset{(-1.59)}{0.007} - \underset{(-2.85)}{0.023}\, ec_{t-1} + \underset{(7.11)}{0.033}\, d902 + \underset{(6.74)}{0.032}\, d011 - \underset{(-5.15)}{0.219}\, \Delta(\pi_t) \\ & - \underset{(-2.60)}{0.114}\, \Delta(\pi_{t-1}) + \underset{(2.84)}{0.209}\, \Delta(m-p)_{t-1} + \underset{(1.92)}{0.139}\, \Delta(m-p)_{t-2} + \hat{u}_t \end{aligned}$$

T=88 (1983.1-2004.4)

According to the negative coefficient of the error correction term *ec*, excess money lowers money growth, as one expects in a stable model. In addition, changes in inflation are highly significant. The results point to substantial inertia in the adjustment of real money balances, as the feedback coefficient is very low and two lagged changes of money demand are relevant in the preferred specification. Finally, as the *t*-values indicate, the impulse dummies *d*902 and *d*011 should enter a money demand equation.

Standard specification tests are largely supportive for the model, see table 3. *LM* is a Lagrange Multiplier test for autocorrelation in the residuals up to order 1, 4 and 8. The *p*-values show, that no problems with autocorrelated residuals occur. *ARCH* is a Lagrange multiplier test for conditional heteroskedasticity. Again, the residuals do not exhibit such kind of behaviour. Furthermore, they are distributed as normal, as indicated by the Jarque-Bera test. The cusums of squares test does not indicate any structural break in the regression coefficients, see figure 4. However, the Ramsey *RESET* test might point to some nonlinearities in the relationships, as the third order power of the fitted endogenous variable turns out to be significant at the 0.05 level (RESET2).

-Table 3 and figure 4 about here-

Because of the ambiguous results of the *RESET* procedure, tests on the functional form are performed as well. The idea is, that the linear model (4) can be possibly further improved, if nonlinearities are taken into account. In fact, Carstensen (2004) and several other autors (e.g. Chen and Wu, 2005 for the US and the UK) have emphasized the presence of nonlinearities in money demand behaviour. Lütkepohl, Teräsvirta and Wolters (1999) have detected nonlinearities in the development of the German M1 aggregate. In this study, smooth transition regression (STR) techniques are used to discriminate between linear and nonlinear alternatives, see Granger and Teräsvirta (1993) and Teräsvirta (2004). Due to this choice, two benefits can be exploited. First, the methods provide a convenient framework to examine the null hypothesis of linearity. Second, in case the the null is rejected, a STR model will govern the nonlinear relationship between money demand and the explanatory variables. As the STR approach allows either for smooth or sudden changes in the parameter regime over time, the tests can also be viewed as tools to uncover non-constancy in the regression parameters. More precisely, the STR model takes the form

(5) $$z_t = \beta' w_t + (\pi_1 + \pi_2 G(h_t, \alpha))' x_t + u_t$$

where z is the dependent variable, w and x are subvectors of regressors, u the error term with white noise properties and G a continuous transition function customarily bounded between 0 and 1, whose parameters are denoted by α. The transition function G depends on the transition variable h, which may include elements of w and x and causes the coefficient vector $\pi_1+\pi_2G(h_t,\alpha)$ to be non-constant. For example, if $h_t=t$, the vector $\pi_1+\pi_2G$ changes smoothly over time. But, the transition variable might also be stochastic, and each of the regressors in (4) is considered as a possible candidate.

If G=0, the model is linear, and this hypothesis can be tested after specifying the respective form of the transition function G. Both logistic and exponential specifications are considered. It should be noted, however, that the alternative is not identified under the null. Therefore, an auxillary regression

$$(6) \qquad z_t = \beta' w_t + \delta_0' x_t + \delta_1' x_t h_t + \delta_2' x_t h_t^2 + \delta_3' x_t h_t^3$$

is needed. The null of linearity is tested via the hypothesis $\delta_1=\delta_2=\delta_3=0$, and this test has power against logistic and exponential alternatives, see Granger and Teräsvirta (1993). To investigate the issue of parameter constancy, three tests are available. Under the alternative, the parameters may change monotonically over time (LM1), symmetrically with respect to an unknown point in time (LM2) or non-monotonically but in a non-symmetric way (LM3). The results of the linearity tests are exibited in table 4, and those from the parameter constancy tests in table 5. Computations are carried out by using JMulTi 4.05, see Teräsvirta (2004).

-Tables 4 and 5 about here-

If the autoregressive terms in (4) are taken as transition variables, the test statistic for the linearity test cannot be computed because of near singularity of the moment matrix in the auxillary regression (6). Otherwise, the tests point to the linear specification, as the null is not rejected in any case. Numerical problems are also apparent in the tests on parameter constancy. Hence, tesing is done under two specifications, where either the first or the second lagged endogeneous variable is excluded from consideration. As the

LM tests do not provide any evidence against the null of parameter constancy, the linear specification is strongly preferred.

6 Conclusion

In this paper we have analysed money demand behaviour in the euro area, where special emphasis is given to the issue of stability. In fact, many researchers have detected instabilities especially when data after 2001 are included in the analysis. Such a result casts serious doubts concerning the rationale of monetary aggregates in the monetary strategy of the ECB. In contrast to the bulk of the literature, we report strong evidence in favour of a stable money demand relationship. This result can be achieved by an appropriate interpretation of the role of inflation in the cointegration vector. If the analysis is done without imposing a short run homogeneity restriction between money and prices, a long run money demand relationship is identified, where recursive estimation lead to stable long run parameters. In addition, the corresponding error correction model is robust to a wide array of specification tests, including procedures for nonlinearities and parameter instability. Hence, the apparent monetary overhang is in line with standard models of money demand behaviour, and is not expected to lead to a rise in inflation.

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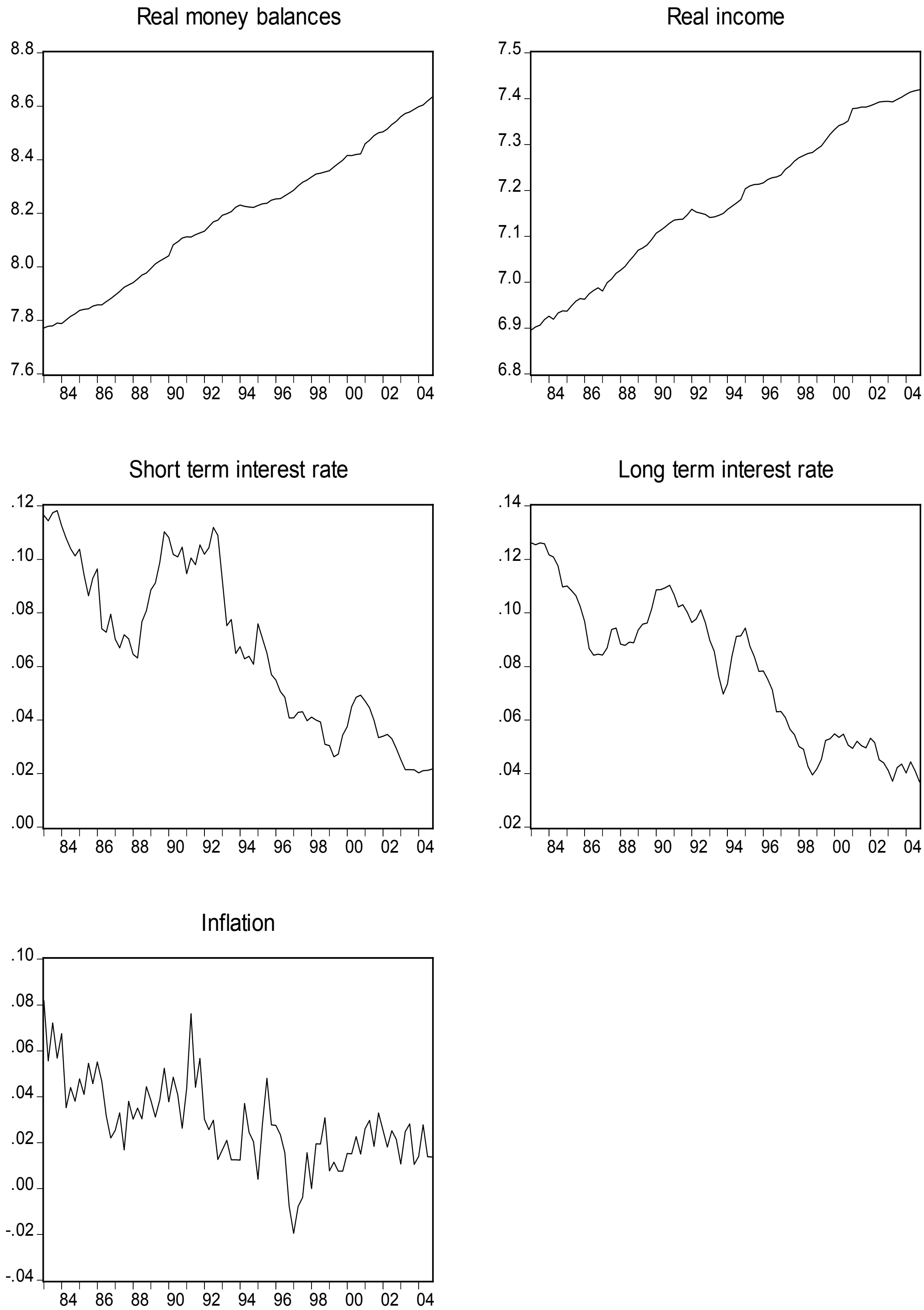
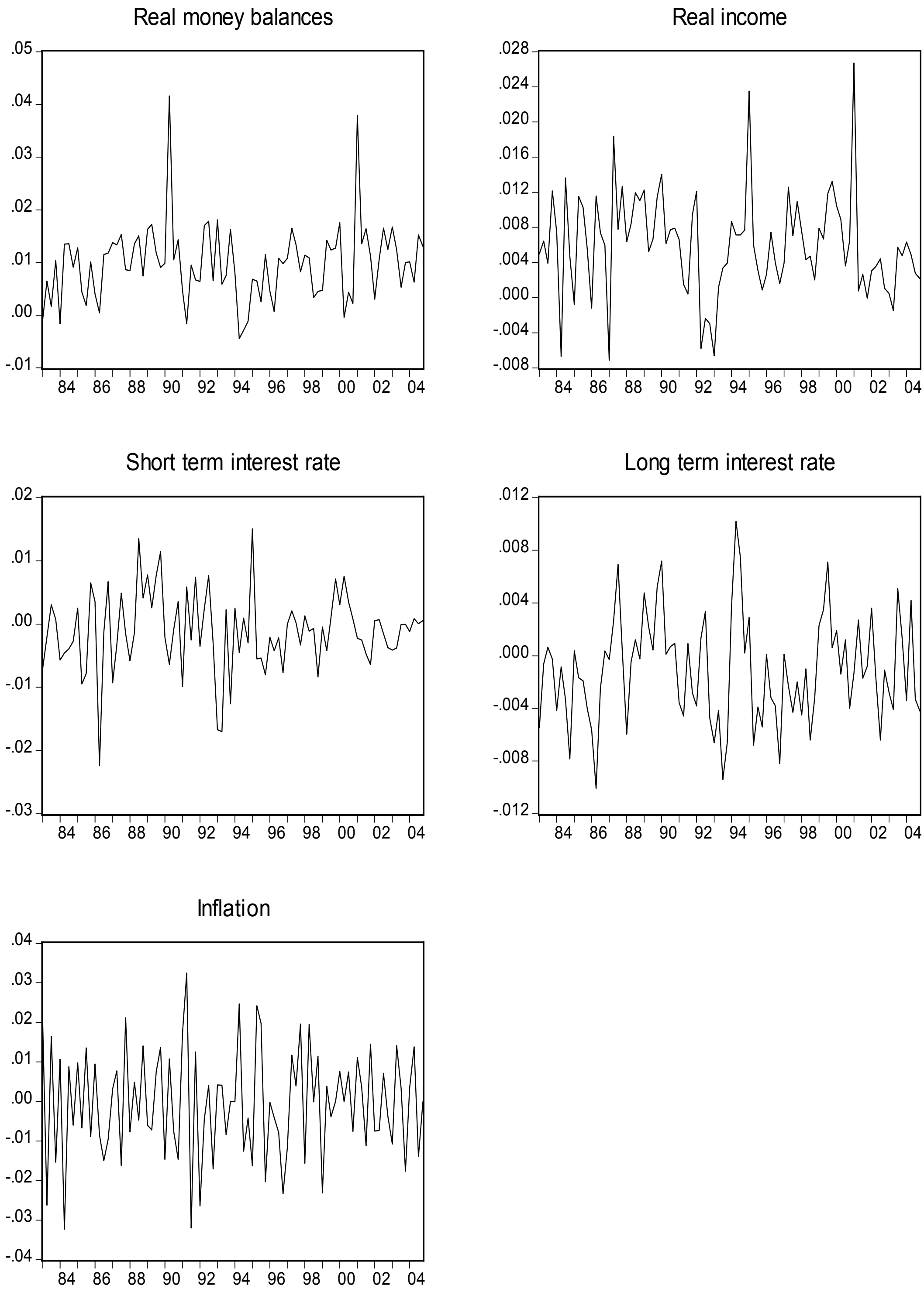
Figure 1: Variables used in the empirical analysis (levels)



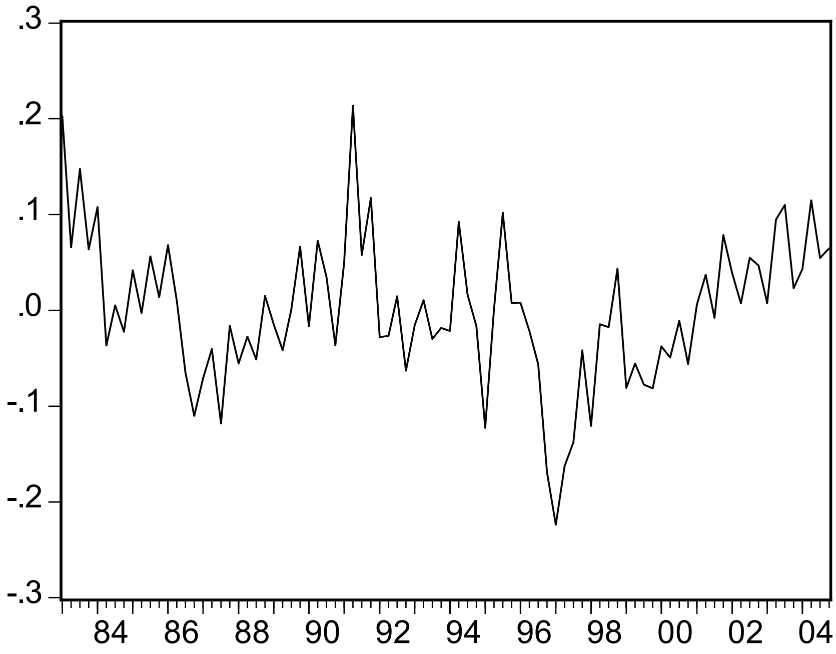
Note: Sample period 1983.1-2004.4. Real money and real GDP in logarithms and deflated by the GDP deflator.

Figure 2: Variables used in the empirical analysis (first differences)



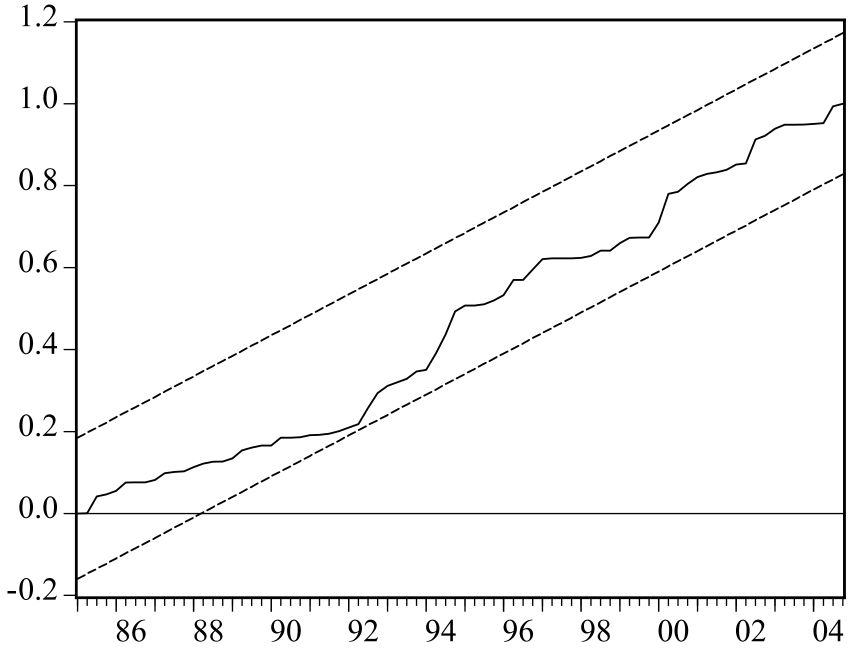
Note: Sample period 1983.1-2004.4. Real money and real GDP in logarithms and deflated by the GDP deflator.

Figure 3: Mean-adjusted deviations from the long run



Note: Sample period 1983.1-2004.4. Long run estimated according to equation (3a).

Figure 4: Cusum of squares from the error correction model



Note: Sample period 1983.1-2004.4. Dashed lines represent 0.05 significance levels.

Table 1: Cointegration tests for sample period 1983.1-2004.4

Variables	Deterministics	Rank null hypothesis	Johansen trace test Lag order	Finite sample correction
m-p, *y*	con_u, dum	0 1	9.04 2.01 [n=1]	
rl, *π*	con_r	0 1	19.82(*) 4.40 [n=2]	18.92(*)
rs, *π*	con_r	0 1	32.10** 4.82 [n=1]	31.37**
rl, *rs*	con_r	0 1	17.35 4.23 [n=2]	
rl, *rs*, *π*	con_r	0 1 2	32.55(*) 17.06(*) 4.49 [n=2]	30.33 15.89
m-p, *y*, *π*	con_u, dum	0 1 2	50.03** 8.39 2.25 [n=1]	48.32**
m-p, *y*, *rl*	con_u, dum	0 1 2	19.12 5.68 1.35 [n=1]	
m-p, *y*, *rs*	con_u, dum	0 1 2	25.57 10.78 0.03 [n=1]	
m-p, *y*, *π*, *rs*	con_u, dum	0 1 2 3	67.56** 24.34 10.59 0.001 [n=1]	64.49**

Note: con_u, con_r = constant unrestricted, restricted, dum = impulse dummies for 1990.2 and 2001.1, respectively. The finite sample correction is due to Reimers (1992). A (*), *, ** denotes significance at the 0.1, 0.05 and 0.01 level. Critical values are from MacKinnon, Haug and Michelis (1999), and are also valid for the finite sample correction. The lag order of the VAR in levels is determined by the Schwarz criterion, with maximum lag order 8. The criteria point to the same lag order in all cases, which is the number in brackets below the test values.

Table 2: Estimated cointegration parameters and trace tests

A Forward recursive estimates

Sample period	$\beta(y)$	$\beta(\pi)$	Trace statistic	Trace statistic corrected
1983.1-1998.4	1.25 (0.06)	-2.92 (0.35)	55.72**	53.10**
1983.1-1999.4	1.26 (0.05)	-2.90 (0.33)	60.32**	57.66**
1983.1-2000.4	1.20 (0.06)	-3.29 (0.37)	57.97**	55.56**
1983.1-2001.4	1.29 (0.05)	-2.80 (0.34)	55.91**	53.70**
1983.1-2002.4	1.32 (0.05)	-2.89 (0.36)	54.47**	52.43**
1983.1-2003.4	1.30 (0.06)	-3.67 (0.49)	52.53**	50.66**
1983.1-2004.4	1.24 (0.09)	-5.16 (0.72)	50.03**	48.32**

B Backward recursive estimates

Sample period	$\beta(y)$	$\beta(\pi)$	Trace statistic	Trace statistic corrected
1983.1-2004.4	1.24 (0.09)	-5.16 (0.72)	50.03**	48.32**
1984.1-2004.4	1.28 (0.08)	-5.02 (0.70)	46.21**	44.56**
1985.1-2004.4	1.30 (0.08)	-4.67 (0.70)	42.76**	41.16**
1986.1-2004.4	1.33 (0.08)	-4.61 (0.71)	40.46**	38.86**
1987.1-2004.4	1.28 (0.09)	-4.59 (0.75)	38.08**	36.50**
1988.1-2004.4	1.30 (0.10)	-4.27 (0.72)	36.21**	34.61*
1989.1-2004.4	1.34 (0.12)	-4.48 (0.78)	32.24*	30.73*
1990.1-2004.4	1.44 (0.16)	-5.95 (0.98)	33.59*	31.91*

1991.1-2004.4	1.54 (0.18)	-7.33 (1.15)	35.62*	33.71*

C Moving window (64 observations)

Sample period	$\beta(y)$	$\beta(\pi)$	Trace statistic	Trace statistic corrected
1983.1-1998.4	1.25 (0.06)	-2.92 (0.35)	55.72**	53.10**
1984.1-1999.4	1.29 (0.05)	-2.85 (0.33)	55.31**	52.72**
1985.1-2000.4	1.24 (0.06)	-3.10 (0.38)	48.95**	46.66**
1986.1-2001.4	1.32 (0.05)	-2.69 (0.36)	44.60**	42.51**
1987.1-2002.4	1.32 (0.05)	-2.61 (0.38)	43.22**	41.20**
1988.1-2003.4	1.31 (0.07)	-3.16 (0.50)	38.80**	36.98**
1989.1-2004.4	1.34 (0.12)	-4.48 (0.78)	32.24*	30.73*

Note: The constant is unrestricted. The finite sample correction is due to Reimers (1992). A (*), *, ** denotes significance at the 0.1, 0.05 and 0.01 level. Critical values are from MacKinnon, Haug and Michelis (1999), and are also valid for the finite sample correction. According to the Schwarz information criterion, no lagged differences are included. Two impulse dummies for 1990.2 and 2001.1 enter the models, if the sample period considered covers the respective observations.

Table 3: Standard specification tests for the error correction equation

R2=0.59	SE=0.005	JB=1.46 (0.48)
LM(1)=0.06 (0.81)	LM(4)=1.80 (0.14)	LM(8)=1.32 (0.25)
ARCH(1)=1.71 (0.19)	ARCH(4)=1.26 (0.29)	ARCH(8)=0.63(0.75)
RESET(1)=2.52 (0.12)	RESET(2)=3.25 (0.04)	RESET(3)=2.32(0.08)

Note: Sample period 1983.1-2004.4. R2=R squared adjusted, SE=standard error of regression, JB=Jarque-Bera test, LM=Lagrange multiplier test for no autocorrelation in the residuals, ARCH=Lagrange multiplier test against conditional heteroscedasticity, RESET=Ramsey test, *p*-values in parantheses.

Table 4: *p*-values of linearity tests of error correction model (4) against smooth transition alternatives for different transition variables

Transition variable	ec_{t-1}	$\Delta(\pi_t)$	$\Delta(\pi_{t-1})$	$\Delta(m\text{-}p)_{t-1}$	$\Delta(m\text{-}p)_{t-2}$	t
p-value	0.73	0.16	0.92	NA	NA	0.51

Note: Sample period 1983.1-2004.4. All tests are F-tests, based on the auxillary regression (6), as discussed in Granger and Teräsvirta (1993). NA=Test statistic cannot be computed due to invertibility problems in the auxillary regression.

Table 5: *p*-values of various LM-type tests on parameter constancy

A Coefficient of $\Delta(m\text{-}p)_{t-1}$ excluded

LM1	LM2	LM3
1.19 (0.32)	0.76 (0.75)	1.00 (0.43)

B Coefficient of $\Delta(m\text{-}p)_{t-2}$ excluded

LM1	LM2	LM3
1.32 (0.23)	0.82 (0.69)	0.91 (0.60)

Note: Sample period 1983.1-2004.4. *F*-tests for parameter constancy, *p*-values in parantheses. Due to matrix inversion problems, the first (second) lagged endogeneous variable is excluded from the constancy test in the upper (lower) part of the table.