MODELLING MONEY DEMAND IN GERMANY

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SUMMARY

In this paper an empirically stable money demand model for M3 in Germany is presented. The sample period 1975–94 includes German unification. It is shown that this development has not substantially destabilized money demand. Parameter stability is extensively tested and not rejected. Applying encompassing tests, this model encompasses two recent models but is not encompassed by them. Exogeneity of the explanatory variables is discussed and tested along the definitions given in Engle, Hendry and Richard (1983). There is evidence that inflation and long-term interest rates are super-exogenous with respect to the parameters of the demand for M3 model. This result and the empirical long-run money demand function presented in this paper may affect the applicability of the so called ‘P-Star concept’ for German M3.

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1. INTRODUCTION

In this paper I will show that German unification has not substantially destabilized the demand for money M3 in Germany. Using quarterly data from 1975 to the fourth quarter of 1994, I re-estimate my earlier money demand model which is based on data until the end of 1989 (see Beyer, 1994) and which hence does not take into account the 1990 unification. Apart from a few impulse dummy variables — but no step dummy — the model presented here has exactly the same parameterization and it keeps all its properties. Various recently suggested tests for parameter instability which take into account the uncertainty of the date of a break will be applied such as Andrews and Ploberger’s (1994) Mean Wald and Exponential Wald tests. The focus will be put on the exogeneity status of the variables that explain M3. I will show evidence that in Germany inflation and long-term interest rates are super-exogenous with respect to the parameters of a demand function for M3, using the notion of ‘super-exogeneity’, as defined in Engle, Hendry and Richard (1983). This result and the specification of the long-run money demand function presented here are discussed with respect to the Bundesbank’s application of the so-called ‘P-Star concept’ by Hallman et al. (1991) which is used as a measure to show the impact of movements of money stock on inflation. For German data it was recently applied by Toedter and Reimers (1994), who combine the P-Star concept with a long-run money demand function to circumvent the ‘serious weakness that lies on the assumption of a constant velocity of circulation’ (p. 273).

The remainder of the paper is organized as follows. In Section 2 an empirically stable error-correction model for German money M3 is estimated. A cointegration analysis is presented and weak exogeneity is successfully tested within the Johansen procedure, justifying re-estimation of the earlier money demand model in a single-equation framework with an extended data set.

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Section 3 analyses economic and statistical properties of the empirical model. A long-run money demand function is derived and its economic implications are discussed. It is shown that its coefficients did not change substantially compared with those of the pre-unification function in Beyer (1994) and parameter instability is formally tested. Following the strategy of Hendry and Ericsson (1991b) and Engle and Hendry (1993) the variables for inflation and long-term interest rates of the money demand model are tested for super-exogeneity. The model’s ability to explain properties of other models (including seasonality) is tested via encompassing (see e.g. Mizon and Richard, 1986). Section 4 briefly discusses implications of the empirical results obtained in this paper and contrasts them to the Bundesbank approach. Section 5 concludes the main results. The Appendix provides the sources of the data.

2. AN EMPIRICAL MODEL FOR GERMAN M3

Evidence of the empirical stability of German money demand after unification is mixed and a matter of controversy. In this section an empirically stable error-correction model for real M3 is presented. Before estimating the model, tests for unit roots in the relevant variables and a cointegration analysis due to the theoretical long-run formulation of the model are discussed.¹

2.1. Cointegration and the Long-run Theoretical Model

In contrast to the long run money demand function of Toedter/Reimers (1994), Issing/Toedter (1994) and Deutsche Bundesbank (1995), where the coefficients are derived directly from a static cointegration regression, I will derive a long-run relationship from a dynamic empirical econometric model following the modelling strategy of, inter alia, Hendry and Ericsson (1991a,b). Let

\[ (m - p)^* = \delta_0 + \delta_1 y + \delta_2 RS + \delta_3 RL + \delta_4 \Delta_4 p \]  

be the long-run function for the real money stock M3 where \( p \) is represented by the GDP deflator with \( \Delta_4 p \) being the annual inflation rate; \( y \) is real GDP in 1991 prices. Lower-case letters denote here and elsewhere natural logarithms. RS is a short-term interest rate and measures the own rate of M3. It is represented by the Frankfurt three-month funds money market rate. RL is a long-term interest rate represented by yields on outstanding public bonds and hence a measure for the opportunity cost of holding M3. The data are quarterly series from 1975(1) until 1994(4). As a first rough guide the results of DF and ADF tests in Table I suggest that \( m \) and \( p \) are I(2) variables and \( (m - p), y, \Delta_4 p, RS, \) and \( RL \) are I(1) variables. To avoid modelling I(2) variables, this suggests the imposition of long-run price homogeneity such that \( (m - p) \), is the variable of interest as postulated in equation (1). Assuming that long-run velocity of M3 is a function of \( \Delta_4 p \), RS and RL such that

\[ (m - p - y)^* = \delta_0^* + \delta_2^* RS + \delta_3^* RL + \delta_4^* \Delta_4 p \]  

then for equation (2) to be a valid representation of the (inverse) velocity of M3 it is necessary that there exists at least one cointegrating vector \( \beta \) forming a cointegrating relationship \( \beta' x \) where \( x = [(m - p), y, \Delta_4 p, RS, RL] \). The number of cointegrating vectors was tested by applying

¹ All estimates and tests in Section 2 were calculated using PC-GIVE 8.10 (Doornik and Hendry, 1994).
Table I. Augmented Dickey–Fuller statistics

<table>
<thead>
<tr>
<th>Variable</th>
<th>Specification</th>
<th>m</th>
<th>Δm</th>
<th>ΔΔm</th>
<th>p</th>
<th>Δp</th>
<th>ΔΔp</th>
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<table>
<thead>
<tr>
<th>Variable</th>
<th>y</th>
<th>Δy</th>
<th>RS</th>
<th>ΔRS</th>
<th>RL</th>
<th>ΔRL</th>
<th>(m – p)</th>
<th>Δ(m – p)</th>
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<td>Longest lag</td>
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</table>

Based on the assumption that the rank of the long-run matrix \( \beta^r = \Pi \) is unity, it was then tested if \( (y_i, \Delta \Delta p_i, RS_i, RL_i) \) are weakly exogenous for the cointegrating vector. The weighting

\[
(m – p) = 0.94 y – 1.78 \Delta \Delta p + 1.6 RS – 3.28 RL + 0.002 Trend
\]

\(^2\) D90(1) is included because of exceptional movements in interest rates due to inflationary expectations after the opening of the wall in November 1989. D91(1) captures the shift in the series for \( m \) and \( y \) which are recorded until 1990(4) for West Germany only.

\(^3\) Using degrees-of-freedom adjusted critical values, none of the Johansen test statistics is significant at 5%. However, the test results are consistent with Monte Carlo evidence by Kostiel (1994) that in small samples the Johansen procedure underestimates the dimension of the cointegrating space. As in Hendry and Doornik (1994), the outcome of the cointegration tests depends also here on the choice of restriction of the deterministic variables. If D90(1) is restricted to be in the cointegrating space the trace test rejects the null of no cointegration at 1% and the maximum eigenvalue test at 5%, thus favouring the assumption of one cointegrating vector. If, however, D91(1) is restricted instead of D90(1) the trace test even rejects at 5% the null of one in favour of two cointegrating vectors.
The cointegrating relationship enters only in the $m\cdot p$ equation of the system. The last row in Table II shows the test results for separate and joint restrictions on the $a$’s, giving strong evidence that weak exogeneity can be assumed. Two restrictions on $b$ are imposed: the coefficient on income being unity for income homogeneity of real money demand and the trend being zero. Both restrictions are not rejected by the LR test statistic $w^2$, yielding a cointegrating vector $m\cdot p = y - 4.315\Delta_p + 5.184RS - 8.484RL$.

If the restrictions on $x$ and $\beta$ are imposed jointly, the resulting cointegrating vector is

$$ (m - p) = y - 4.315\Delta_p + 5.184RS - 8.484RL $$

and $\chi^2(6) = 13.591 [0.035]$ being significant at 5% but not at 1%.
In summary, the test results of the cointegration analysis suggest that it seems reasonable to proceed with a single-equation analysis for \((m - p)\). Then, the long-run function for \((m - p)\) can be derived from a sequentially reduced unrestricted ADL model. Before doing so, the explanatory variables were tested for strong exogeneity by assuming weak exogeneity and applying Granger-causality tests (four lags, \(T = 1976(2)–1994(4)\)) for feedbacks from \(\Delta(m - p)\). Since \((m - p)\) and the four explanatory variables are cointegrated Granger-causality can also be tested within the complete VAR model in I(1) space. In this case test statistics for excluding coefficients of lagged \((m - p)\), have a limiting \(\chi^2\) distribution as shown, for example, in Watson (1994). The results reported in Table III for the test statistics \(F(4, 65)\) and \(\chi^2(4)\) for individual and \(\chi^2(12)\) and \(\chi^2(16)\) for joint tests show that strong exogeneity is rejected for real GDP when tested individually but is not rejected for each of the other variables. This implies that there is neither a direct influence of real money on the weak exogenous variables nor an indirect feedback on inflation and interest rates via feedbacks from the error-correction term.

### 2.2. A Single-equation Conditional Model for Money Demand

Let \(((m - p), y, \Delta p_t, RS_t, RL_t) = Z_t\) be a process represented by a joint normal distribution and with \(Z_{t-1}^0 = (Z_1, Z_2, \ldots, Z_{t-1})\) representing the history of \(Z_t\). Factorization into conditional and marginal distribution yields

\[
D(Z_t | Z_{t-1}^0; \psi) = D((m - p) | \Delta p_t, y, RS_t, RL_t, Z_{t-1}^0; \psi_1) \\
\cdot D(\Delta p_t, y, RS_t, RL_t | Z_{t-1}^0; \psi_2)
\]

(6)

Weak exogeneity of \((y, \Delta p_t, RS_t, RL_t)\) for the parameters of the conditional distribution is necessary for modelling \(\Delta(m - p)\) within a single-equation framework. Using theorem 8.1 in Johansen (1995) the conditional distribution of \((m - p)\), can then be represented by an error-correction model that explains changes in \((m - p)\), by its own lags, the error-correction term, and by simultaneous changes and their lags of the weak exogenous variables. The model may also contain deterministic terms like a constant and dummies which are represented in the variable \(D_t\). For a single-equation model of \((m - p)\), to be estimable by OLS within a structural VAR framework would imply restricting the coefficients of \((m - p)\), to be zero in the four equations for the four variables of the marginal distribution in equation (6); covariances between structural shocks of money demand and those of each of the four other equations have to be assumed to be zero.
These identifying restrictions are consistent with an economy in which output and prices do not adjust instantaneously and monetary policy is conducted via controlling the (short-term) interest rate. It is assumed that within a quarter money demand is accommodated by money supply. Long-run interest rates, which cannot be controlled directly by the Bundesbank, are assumed not to react instantaneously due to publication lags of money stock figures. Starting from the general error-correction model with four lags which is a reparameterized ADL(5,5) model

\[ \Delta(m - p)_t = \sum_{i=0}^{4} c_i \Delta(m - p)_{t-i} + \sum_{i=0}^{4} d_{1,i} \Delta \Delta p_{t-i} + \sum_{i=0}^{4} d_{2,i} \Delta y_{t-i} + \sum_{i=0}^{4} d_{3,i} \Delta RS_{t-i} + \sum_{i=0}^{4} d_{4,i} \Delta RL_{t-i} + \lambda (m - p - y)_{t-1} + \gamma_1 RS_{t-1} + \gamma_2 RL_{t-1} + \Phi^i D_i + u_t \]  

(7)

Equation (7) is stepwise simplified and reparameterized to the following empirical econometric error-correction model (with standard deviations in parentheses):

\[ \Delta(m - p)_t = 0.09 (0.05) \Delta(m - p)_{t-4} - 0.76 (0.18) \Delta \Delta p_{t} + 0.12 (0.045) \Delta y_{t-2} - 0.32 (0.09) (\Delta RS_{t-1} + \Delta RS_{t-3}) + 0.34 (0.16) \Delta^2 RL_{t} + 0.28 (0.11) RS_{t-1} - 0.58 (0.17) RL_{t-1} - 0.29 (0.16) \Delta p_{t-1} - 0.08 (0.0275) (m - p - y)_{t-1} + 0.10 (0.03) + 0.017 (0.007) D86(1) - 0.015 (0.007) D90(1) + 0.13 (0.007) D91(1) + 0.012 (0.007) D92(3) - 0.019 (0.007) D94(4) + \hat{u}_t \]  

(8)

\[ R^2 = 0.87; F(14, 60) = 29.419 [0.0000]; \sigma = 0.674\% \]

\[ DW = 1.77; \text{RSS} = 0.0027; T = 75 [1976(2)–1994(4)] \]

AR 1–5 F(5, 55) = 0.61 [0.70]

X2 F(28, 31) = 0.31 [0.99]

RESET F(1, 59) = 0.05 [0.82]

ARCH 4 F(4, 52) = 1.14 [0.35]

Normality \[ \chi^2(2) = 1.70 [0.43] \]

Figure 1 shows actual and fitted values of equation (8). The model has the same parameterization and its coefficients are sufficiently close to Beyer’s (1994) pre-unification model, except for the four post-unification dummy variables. Note, however, that no step dummy is needed for ensuring parameter constancy after unification and that D90(1) is the only dummy with an explicit economic motivation due to unification.  

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4 This corresponds with the Deutsche Bundesbank’s (1994) view that it ‘cannot directly limit the expansion of the money stock at will, simply by failing to meet any excess demand for central bank balances on the part of banks nor is it able to offset unduly weak demand for central bank money so smoothly that the increase in the money stock will at no time lag behind the target set’ (p. 122).

5 D86(1) captures the rise in increase of M3 in December and January 1986 (see Deutsche Bundesbank, 1986). D92(3) is set with regard to a huge increase in M3 due to the crisis of the European Currency Mechanism in September 1992 (see Deutsche Bundesbank, 1992b). Imposing D94(4) is motivated by institutional changes. It captures the decrease of M3 due to flows into newly introduced money market funds.
3. THE PROPERTIES OF THE MODEL

3.1. The Long-run Solution

The theoretical parameters of interest $\phi = (\delta_0, \ldots, \delta_4)$ of the theoretical model (1) can now be derived from the estimated model (8). The long-run steady-state solution of the model is based on the annual growth rates $\Delta_4m = g_m$; $\Delta_4p = g_p$; and $\Delta_4y = g_y$ and assuming no growth in the two interest rates:

$$\frac{1}{4}(g_m - g_p) = \frac{0.09}{4}(g_m - g_p) - 0.29g_p + \frac{0.12}{2}g_y + 0.28RS - 0.58RL - 0.08(m - p - y) + 0.10$$

'Solved' for the inverse velocity and under the condition that $g_m = (g_y + g_p)$ this yields equation (2*):

$$(m - p - y) = K^* + 3.5RS - 7.25RL$$

where $K^* = 1.25 - 3.625g_p - 2.09g_y$  

(2*)

which implies the steady-state growth path of real money stock M3 (1*):

$$(m - p)^* = 1.25 - 2.09g_y + y - 3.625\Delta_4p + 3.5RS - 7.25RL$$  

(1*)
It follows from the semi-elasticities of interest rates and inflation rate that in the long run an increase in the 3-month rate $RS$ of 1% is accompanied by an increase in M3 of 3.5% whereas an increase of the long-run interest rate $RL$ of one percentage point is in line with a decrease of real M3 by 7.25%, each $ceteris paribus$. Likewise, an increase in the annual inflation rate of one percentage point is accompanied by a decrease of real M3 of 3.625% and because of unit price homogeneity the latter result is also valid for nominal M3.

3.2. Testing for Parameter Instability

Comparison of equation (1*) with the pre-unification result $\phi = [(1.22 - 3.67 g), 1, -3.67, 3.67, -6.44]$ in Beyer (1994) which is based on the estimation period 1976(2)–1989(4) shows clearly that unification did not have a major impact on the long-run parameters. The same applies for equation (8). As the RLS graphics in Figures 2 and 3 show, the one-step residuals are within the bands of plus or minus twice the corresponding equation standard errors, and the coefficients of the model are stable over time. Chow tests for structural changes within forecast periods between 1991 and 1994 are insignificant.\(^6\) Chow’s ‘predictive-failure’ test (in PCGIVE the so-called breakpoint $F$-Test) is nowhere significant, with $F(5, 58) = 2.00$ [0.09] being the largest.

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\(^6\) 1991(1)–1994(4): $F(16, 44) = 1.33$ [0.22]; 1992(1)–1994(4): $F(12, 48) = 1.35$ [0.22]; 1993(1)–1994(4): $F(8, 52) = 1.34$ [0.25]. Furthermore, none of the $t$-statistics for a zero forecast innovation mean for the respective forecast period is significant at conventional levels.

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Figure 2. Model (8): one-step residuals and plus/minus twice standard errors
The so-called ‘one-step’ Chow test (see Doornik and Hendry, 1994) is significant at 5% only for two sample periods (in 1988-3 $F(1, 34) = 5.09 [0.03]$ and in 1992-3 $F(1, 50) = 4.28 [0.05]$). However, it is well known that these tests imply an a priori selection of a breakpoint whereas the date of structural change is not defined under the null hypothesis. Standard testing theory is not applicable for a sequence of individual Chow tests and hence the usual $F$-critical values cannot be used. Therefore these tests provide only informal evidence about (in-)stability but might help in improving the ‘design’ of a well-specified model. Another commonly used test is the CUSUM test by Brown, Durbin and Evans (1975). For equation (8) none of the test statistics is significant at 5%. However, in Monte Carlo studies by Andrews, Lee and Ploberger (1996) the CUSUM test performs very poorly and is outperformed by alternative tests for structural breaks. Four alternative testing procedures for which the uncertainty of the break date is explicitly accounted for in the choice of the critical values were used to test equation (8) and the long-run relationship (1) for parameter instability:

1. Quandt’s likelihood ratio (QLR) or ‘SupF’ test;
2. The Mean Wald (MW) or ‘MeanF’ test;
3. The Average Exponential Wald (EW) test;

The results are listed in Tables IV(a) and IV(b).

Figure 3. Model (8): recursive graphs of coefficients
Table IV(a). Tests for structural break

<table>
<thead>
<tr>
<th>Model</th>
<th>Statistic:</th>
<th>QLR¹</th>
<th>MW²</th>
<th>EW³</th>
<th>$L_C^4$</th>
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<tbody>
<tr>
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<td>Method:</td>
<td>FMOLS</td>
<td>OLS</td>
<td>FMOLS</td>
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<tr>
<td>$(m - p)_{(1)}$</td>
<td>Test stat.</td>
<td>18.06</td>
<td>7.46</td>
<td>0.75</td>
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<tr>
<td>Cointegration $p$-value</td>
<td>0.08</td>
<td>0.13</td>
<td>0.10</td>
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<td>I(1) space 5% c.v.</td>
<td>19.08</td>
<td>9.08</td>
<td>0.93</td>
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<tr>
<td>$\Delta(m - p)_{(8')}$</td>
<td>Test stat.</td>
<td>24.82</td>
<td>7.50</td>
<td>5.35</td>
<td>2.90</td>
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<tr>
<td>I(0) space $p$-value</td>
<td>[0.1-10]</td>
<td>17.85</td>
<td>11.55</td>
<td>3.15</td>
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<tr>
<td>$\Delta \Delta d_{p_t}$ (9)</td>
<td>Test stat.</td>
<td>8.99</td>
<td>1.93</td>
<td>1.01</td>
<td>1.14</td>
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<td>marginal model $p$-value</td>
<td>[0.1-10]</td>
<td>7.67</td>
<td>5.23</td>
<td>1.36</td>
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<tr>
<td>I(0) space 5% c.v.</td>
<td>9.49</td>
<td>7.67</td>
<td>5.23</td>
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<tr>
<td>$\Delta RL_t$ (10)</td>
<td>Test stat.</td>
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<td>0.7</td>
<td>0.4</td>
<td>0.34</td>
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<tr>
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<td>0.7</td>
<td>0.4</td>
<td>0.34</td>
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<tr>
<td>I(0) space 5% c.v.</td>
<td>7.81</td>
<td>6.07</td>
<td>4.42</td>
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</tbody>
</table>

¹ Quandt (1960) likelihood ratio (‘SupF’) statistic:

$$QLR = \sup_{\delta \in \delta_0, \delta_1} F_T(\delta)$$

² Hansen (1992a), Andrews and Ploberger (1994) Mean Wald (‘MeanF’) statistic:

$$MW = \int_{\delta_0}^{\delta_1} F_T(\delta) \, d\delta$$

³ Andrews and Ploberger (1994) average exponential Wald test:

$$EW = \ln \left( \int_{\delta_0}^{\delta_1} \exp \left( \frac{1}{2} F_T(\delta) \right) \, d\delta \right)$$

where $F_T(\delta)$ is a sequence of Wald test statistics and $\delta = t/T$ for $t = t_0, \ldots, t_1$; $\delta_0, \delta_1$ are upper and lower bounds of the trimming region (15% for models (1), (9), (10) and 17% for model (8')).

⁴ Hansen (1992a) $L_C$ statistic (FMOLS):

$$L_C = \text{tr} \left\{ M_m^{-1} \sum_{i=1}^{n} S_i \hat{\Omega}_{i:1:2}^{-1} S_i^t \right\}$$

$u_i = y_i - \beta \beta_i$ are the residuals of the cointegrating equation with stochastic regressors $x_i$ and residuals $u_t$ of their DGP; $\hat{u}_i = (u_i, n_{u_i})$ has zero mean and $\Omega_{i:1} = \Omega_{12} = \Omega_{12} = \Omega_{22} = \Omega_{21}$ is the long-run variance of $u_t$ conditional on $u_{2t}$ with

$$\Omega = \lim_{n \to \infty} \frac{1}{n} \sum_{i=1}^{n} \sum_{j=1}^{n} \mathbb{E}(u_i u_i)$$

and $S_i$ contains the cumulative FM first-order conditions.

Hansen (1992b) $L_C$ statistic (OLS):

$$L_C = \frac{1}{n} \sum_{i=1}^{n} S_i V^{-1} S_i$$

where, respectively, $S_i$ contains cumulative and $V$ the squared OLS first-order conditions from the regression $y_i = \beta \beta_i x_i + \chi_i$.


Table IV(b). Model (8') and Hansen's (1992b) test for parameter instability

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Const.</th>
<th>$\Delta(m - p)_{(1)}$</th>
<th>$\Delta \Delta d_{p_t}$</th>
<th>$\Delta y_{1:2}$</th>
<th>$\Delta RS_{t:1}^{-1}$</th>
<th>$\Delta^2 RL_t$</th>
<th>ECM_{r-1}</th>
<th>$\sigma$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Coefficient</td>
<td>-0.001</td>
<td>0.09</td>
<td>-0.76</td>
<td>0.12</td>
<td>-0.31</td>
<td>0.34</td>
<td>-0.08</td>
<td>0.66%</td>
</tr>
<tr>
<td>$L_1$ test statistic</td>
<td>0.08</td>
<td>0.03</td>
<td>0.07</td>
<td>0.04</td>
<td>0.12</td>
<td>0.10</td>
<td>0.11</td>
<td>0.25</td>
</tr>
</tbody>
</table>
Since the asymptotic distributions of these test statistics depend on the stochastic process describing the regressors, different critical values apply depending whether models contain non-trending or trending regressors (see e.g. Stock, 1994, for a discussion). For I(1) processes Hansen (1992a) derives test statistics and provides asymptotic critical values for QLR, MW, and $L_C$ tests. For non-trending processes, see Andrews (1993) for QLR, Andrews and Ploberger (1994) for MW and EW tests, and Hansen (1992b) for the $L_C$ test. To ensure that the error-correction model contains only non-trending regressors the I(1) regressors in equation (8) have been replaced by the error-correction term $ECM_{t-1}$ which is defined as derivations of actual $(m - p)_t$ from the cointegrating long-run relationship $(1*)$. For the reparameterized model (8') Figure 4 shows the sequence of $F(k, T - 2k)$ statistics for Chow’s (1960) sample split test in LM-form, together with its 5% critical value of Andrews (1993) and the one of the standard $\chi^2(12)$ distribution. With 12 regressors for $T = 75$ observations the chosen symmetric trimming region for the QLR, MW and EW tests is $[0.17; 0.83]$ from observation 13 (1979.3) to 62 (1991.3) and hence including German unification. None of the three test statistics (24.82; 7.5; 5.35) is significant at 10%. The results of Hansen’s (1992b) test for parameter instability of individual parameters of model (8') are reported in Table IV(b). None of the $L_1$ test statistics is significant at

---

7 See Andrews, Lee and Ploberger (1996) for a Monte Carlo analysis comparing the power of these tests and Stock and Watson (1996) for a recent application on US macroeconomic time series.
10%. Model (1) has been estimated by Phillips and Hansen’s (1990) Fully Modified OLS method using the GAUSS application COINT 2.0 by Ouliaris and Phillips (1994). For the estimation of the covariance parameters a quadratic spectral kernel as recommended by Andrews (1991) was used and the implemented trimming region was [0.15, 0.85]. None of the test statistics is significant at 5% with $\sup F = 18.07$ [0.08], $\text{Mean } F = 7.46$ [0.13], and $L_C = 0.75$ [0.10]. Since, as pointed out by Hansen (1992a), $L_C$ is testing the null of cointegration against the alternative of no cointegration this test result is in line with the results of the Johansen test procedure in Section 2.

3.3. Super-exogeneity

For testing super-exogeneity of the contemporaneous weakly exogenous variables $\Delta dP_t$ and $RL_t$ with respect to the parameters of interest of model (8), I followed the strategy that is described in detail in Ericsson (1992) and Engle and Hendry (1993). A test for super-exogeneity is not rejected if one or more shifts in parameters of a marginal model for a weak exogenous variable do not affect the stability of the conditional model (8). The results for tests of parameter instability are reported in Table IV(a). For autoregressive models of $\Delta \Delta dP_t$ (9) and $\Delta RL_t$ (10) only the QLR test for (9) is significant at 10%.

$$
\Delta \Delta dP_t = -0.15 \times 11 \Delta dP_{t-1} + 0.14 \times 11 \Delta dP_{t-2} + 0.18 \times 11 \Delta dP_{t-3} - 0.33 \times 11 \Delta dP_{t-4} + \hat{u}_t \tag{9}
$$

$$
\Delta RL_t = 0.37 \times 12 \Delta RL_{t-1} - 0.14 \times 12 \Delta RL_{t-2} + 0.18 \times 12 \Delta RL_{t-3} + \hat{u}_t \tag{10}
$$

Applying Doornik and Hendry’s (1994) versions of Chow tests, for model (9) neither the breakpoint nor the forecast Chow test statistic is significant at 5% through the estimation period. However, the one-step Chow test is significant at 5% for the periods 1987(3) and 1991(2) and impulse dummies for these periods are significant when they are added as regressors in model (9):

$$
\Delta \Delta dP_t = -0.15 \times 10 \times 10 \Delta dP_{t-1} + 0.18 \times 10 \times 10 \Delta dP_{t-2} + 0.14 \times 10 \times 10 \Delta dP_{t-3} - 0.37 \times 10 \times 10 \Delta dP_{t-4} - 0.01 \times 0.05 \times 87(3) + 0.01 \times 0.05 \times 91(2) + \hat{u}(p) \tag{9'}
$$

As expected, the impulse dummy coefficients are unstable but even the overall statistic for joint instability of all parameters (including the variance) is just significant at only 10% ($L_{13} = 2.90$).

---

8 As expected, the impulse dummy coefficients are unstable but even the overall statistic for joint instability of all parameters (including the variance) is just significant at only 10% ($L_{13} = 2.90$).
For model (10) the forecast Chow test statistics are significant at the beginning of the 1980s — at that time the yield curve of interest rates in Germany was inverted — and the one-step Chow test is significant at 5% for the observations 1980(3) and 1981(4). Inclusion of the respective impulse dummies yields the empirical stable marginal model for $RL (10')$:

$$
\Delta RL_t = 0.37 (0.10)\Delta RL_{t-1} + 0.26 (0.11)\Delta RL_{t-3} - 0.01 (0.004)D80(3) - 0.02 (0.004)D81(4) + \hat{u}(RL),
$$

$R^2 = 0.30; F(4, 70) = 7.3 [0.0001]; \sigma = 0.43\%; DW = 1.87; RSS = 0.0013$; $T = 75 [1976(2)–1994(4)]$

AR 1–4 $F(5, 65) = 1.39 [0.24]$  
$Xi^2 F(8, 61) = 0.37 [0.94]$  
RESET $F(1, 69) = 0.35 [0.56]$  
ARCH 4 $F(4, 62) = 0.55 [0.70]$  
Normality $\chi^2(2) = 1.36 [0.51]$  

However, if the four dummies are added into the money demand model (8), the significance of these additional regressors is individually and also jointly rejected by $F(4, 57) = 0.74 [0.57]$. Applying the Wu–Hausman test by adding the residuals of the marginal models (9') and (10') into (8), their significance is also rejected by $F(2.59) = 0.95 [0.39]$. Thus, inflation and long-term interest rates can be assumed to be super-exogenous for the parameters of interest in model (8). Furthermore, since the coefficients of these variables in the estimated money demand model are empirically stable over time, the null hypothesis that the Lucas Critique does not apply to model (8) was not rejected.

### 3.4. Seasonality and Encompassing

The empirical model (8) and the discussion of weak and strong exogeneity in the previous section is based on seasonally adjusted (SA) data. As shown in Ericsson, Hendry and Tran (1994) (henceforth EHT) even if the cointegrating vector for a set of variables is invariant to the choice of adjusted or non-seasonally adjusted (NSA) data, the exogeneity status may not be. Therefore, a model based on NSA series of M3, GDP and GDP deflator (their logs denoted respectively as $m^u$, $y^u$ and $p^u$) was developed. It was then tested and shown that the SA model (8) appears to encompass the NSA model and two other recently published models.

**Modelling NSA data**

The modelling strategy for the NSA model was essentially the same as for the SA model. First, ADF tests yielded the same provisional conclusions about the order of integration for the NSA variables compared to their SA counterparts. Next, the Johansen procedure for testing the number of cointegrating vectors for the system $[(m^u - p^u)_t, y^u_t, \Delta p^u_t, RS, RL]$ was conducted for a second-order VAR with constant, seasonals and three impulse dummies D90(2), D90(4) and D92(1) each unrestricted, and a restricted deterministic trend. The null of no cointegration ($r = 0$)

---

9 The statistical break in the NSA series due to unification differs with respect to the SA series: $m^u$ is recorded for unified Germany from 1990(2) and $y^u$ from 1992(1) onwards.

10 The results for the ADF tests and cointegration analysis for NSA do not differ substantially from those of SA data, and are not reported in full detail here.
was rejected with a trace statistic of 88.08 at the 5% level. Assuming $r = 1$, the outcomes of the multivariate stationarity test clearly indicate no stationarity for the NSA variables. Weak exogeneity of $(y_t^u, \Delta p_t^u, RS_t, RL_t)$ for the cointegrating vector is clearly not rejected for each individual variable but — in contrast to the SA model — is rejected when jointly tested (see Table V):

Table V. Test statistics for weak exogeneity (NSA data)

<table>
<thead>
<tr>
<th>Restricted $x$</th>
<th>$y^u$</th>
<th>$\Delta p^u$</th>
<th>$RS$</th>
<th>$RL$</th>
<th>Joint $\chi^2(4)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\chi^2(1)$ [p-value]</td>
<td>0.36 [0.55]</td>
<td>2.31 [0.13]</td>
<td>0.03 [0.86]</td>
<td>0.07 [0.76]</td>
<td>14.24 [0.006]</td>
</tr>
</tbody>
</table>

Restricting jointly the coefficients on $y^u$ to unity and trend to zero, and the $x$’s to zero yields a cointegrating relationship (11) that does not differ massively from model (5):

$$(m^u - p^u) = y^u - 2.23 \Delta p^u + 4.852 RS - 11.66 RL$$  (11)

With $\chi^2(6) = 14.804$ [0.022] the imposed restrictions are likewise not rejected at only 1%. Given that weak exogeneity was not rejected individually, the following empirically stable single-equation model (12) for $(m^u - p^u)_t$ was developed:

$$
\Delta(m^u - p^u)_t = -0.65 (0.10)(\Delta p^u_t + \Delta p^u_{t-1}) + 0.13 (0.04)(\Delta y^u_{t-2} + \Delta y^u_{t-3}) \\
- 0.25 (0.12)RS^u_{t-3} - 0.73 (0.17)R L^u_{t-1} \\
+ 0.14 (0.07)RS^u_{t-1} - 0.40 (0.12)RL^u_{t-1} - 0.02 (0.01)(m^u_t - p^u_t - y^u_t)_{t-1} \\
+ 0.09 (0.01) - 0.09 (0.01)S_1 - 0.07 (0.01)S_2 - 0.03 (0.002)S_3 \\
+ 0.018 (0.007)D 8 5(4) + 0.15 (0.007)D 9 0(2) - 0.04 (0.008)D 9 2(4) \\
- 0.03 (0.007)D 9 4(4) + \hat{u}_t
$$  (12)

$R^2 = 0.94; F(14, 60) = 66.409$ [0.000]; $\sigma = 0.656\%$; $DW = 1.89$; $RSS = 0.0026$; $T = 75$ [1976(2)–1994(4)]; AR 1–5 $F(5, 55) = 1.44$ [0.23]; $X^2$ $F(25, 34) = 1.13$ [0.37]; RESET $F(1, 59) = 3.68$ [0.06]; ARCH 4 $F(4, 52) = 2.86$ [0.032]; Normality $\chi^2(2) = 2.06$ [0.36].

The forecast performance of the NSA model for the period 1991(1)–1994(4) is not notably inferior to model (8). None of the $t$-values for a zero forecast innovation mean is significant at conventional levels. However, two Chow test $F$-statistics for the last eight and four quarters, respectively, are significant at 5% ($F(4, 56) = 3.19$ [0.02] and $F(8, 52) = 2.2681$ [0.037]). The long-run money demand function $(1^{*u})$ derived from model (12) for seasonal unadjusted real M3 is

$$(m^u - p^u) = 4.5 - 9.25 g_y + y + 6.5 RS - 20 RL$$  (1$^{*u}$)

Note that $\Delta p^u$ does not enter model $(1^{*u})$. Restricting accordingly the long-run solution (11) within the Johansen framework yields an LR test statistic $\chi^2(7) = 15.29$ [0.03] for joint zero restrictions on $x$, and unity for $y^u$ and a zero trend.
The NSA model has the same number of parameters but a slightly smaller equation standard error than the SA model. Nevertheless, the SA model is able to encompass the NSA model. Conducting an encompassing test as proposed in EHT (1994) by adding the seasonal component \( \{\Delta(m - p), \Delta(m^u - p^u)\} \) and the regressors from the NSA Model (12) to the SA model (8) yields the \( F \)-statistic \( F(12,49) = 1.80 \) [0.07] for the significance of those additional regressors and \( F(11,50) = 1.39 \) [0.21] if the seasonal component is not included. Thus, from the comparative analysis of this section it can be concluded that the NSA model (12) cannot outperform the SA model (8), and given that seasonally adjusted M3 is the relevant target for the Bundesbank’s monetary policy, model (8) is preferred as a reference for examining policy implications.

**Encompassing alternative models**

Two recent models for German M3 that argue in favour of a stable money demand despite unification are those by Kole and Meade (KM, 1995) and by the Bundesbank (Bbk, Deutsche Bundesbank, 1995, p. 32, 4th column). The two main differences of these models compared to model (8) are: their inclusion of wealth in the set of explanatory variables (in real terms as \((wH - p^u)\), in KM and nominal as \((w - yn)\), in the Bundesbank model, where \(yn\) is nominal GDP); the exclusion of a measure for the own rate of interest in M3. Furthermore, the Bundesbank model does not contain any price variable either in the long run or in the short run and no short-run dynamics in the long-term interest rate are included. KM estimate a model for real M3 based on NSA data for the period 1970(3)–1994(4) whereas the Bundesbank model is based on SA data for nominal M3 for the period 1975(1)–1995(1). To test whether model (8) is able to encompass any of these two alternative models they were each re-estimated for the period 1976(2)–1994(4).\(^{12}\) The model standard errors are \(\sigma_{KH} = 0.868\%\) and \(\sigma_{Bbk} = 0.789\%\). Note that models (KH) and (Bbk) are variance dominated by model (8):

\[
\Delta(m^u - p^u)_t = 0.10 \Delta RL_t - 0.93(0.25)RL_{t-1} - 0.94 (0.23)\Delta p_t^u - 0.02 (0.05)(m^u - p^u)_{t-1} \\
+ 0.06 (0.05)y^u_{t-1} - 0.22 (0.09)RL_{t-1} - 0.03 (0.03)(wH - p^u)_{t-1} + 0.11 (0.10) \\
- 0.09 (0.01)S_1 - 0.05 (0.01)S_2 - 0.05 (0.01)S_3 + 0.01 (0.006)US_t \\
- 0.002 (0.007)US_2 + 0.01 (0.006)US_3 + 0.15 (0.01)D90(2) + 0.006 (0.01)D90(3) \\
+ 0.005 (0.01)D92(1) + \hat{u}_t \quad (KH)
\]

\[
\Delta m_t = 0.14 (0.08)\Delta m_{t-1} + 0.45 (0.18)\Delta yn_t + 0.01 (0.17)\Delta yn_{t-1} + 0.4 (0.15)\Delta (w - yn)_t \\
+ 0.24 (0.15)\Delta (w - yn)_{t-1} - 2.24 (0.06)ECT_{t-1} \\
+ 0.11 (0.01)D91(1) + 0.003 (0.003)D91(1s) - 0.01 (0.01)D94(3) \\
- 0.02 (0.01)D94(4) + \hat{u} \quad (Bbk)
\]

where

\[
ECT_t = 0.20 (0.07) - 0.53 (0.18)RL_t + 1.05 (0.02)yn_t + 0.22 (0.03)(w - yn)_t + \hat{\epsilon}_t
\]

Applying the EHT test strategy as above, the null hypothesis that model (8) encompasses the NSA model (KH) is neither rejected when the seasonal component is excluded \(F(13,47) = \)

---

\(^{12}\) Note that the estimates for the smaller sample period here do not match exactly those of their original sample periods.
1-37 [0-14]) nor when it is included ($F(14, 46) = 1.59 [0-12]$). As discussed in EHT, reversing the roles of the SA and NSA model for testing Encompassing is problematic. However, doing so yields strong rejection of excluding the regressors of model (8) from the general model, no matter if the seasonal component is included or not, with $F(13-46) = 4-90 [0-000]$ and $F(12-47) = 3.23 [0-002]$, respectively. When testing that model (8) encompasses (Bbk)—which requires reparameterizing model (8) by adding $D_{pt}$ on both sides—neither the Sargan test statistic $w^2_8 = 10.9 [0-21]$ nor the conventional $F$-test $F(8, 51) = 1.44 [0-2017]$ is significant. Conversely, the null that (Bbk) encompasses model (8) is strongly rejected by $w^2_{14} = 35.1 [0-0001]$ and $F(14, 51) = 4.28 [0-0001]$.

Thus, there is very strong evidence that model (8) encompasses the models of Kole and Meade (1995) and Deutsche Bundesbank (1995), whereas these models appear not to encompass model (8). Note the insignificance of $w^{yn}_{yt}$ when it is included together with five lags in the general ADL(5,5) model that leads to model (8): the $F$-statistic is $F(6, 39) = 1.158 [0.35]$. The coefficient of $(w - yn)$ of the implied static long solution is 0-11 with a standard error of 0-21. The same applies for the NSA variable of wealth ($wH_{yt}$) with respect to model (12). The $F$-statistic for including it in the respective ADL model is $F(6, 33) = 1.086 [0.39]$; the coefficient of $(wH - p^n)$ of the implied static long-run solution is 0-29 with a standard error of 0-44.

4. THE BUNDESBANK POLICY AND THE P-STAR CONCEPT

A paramount role in the theoretical foundation of the Bundesbank’s monetary policy, of which ultimate target is price level stability, is played by an empirically stable money demand function. A stable money demand function is necessary to estimate the equilibrium velocity of the money stock ($v^*$). Together with an estimate of the production potential ($y^*$) the growth of the money stock is derived from the quantity equation

$$m \equiv p^* + y^* - v^*$$

$$\Delta m \equiv \Delta p^* + \Delta y^* - \Delta v^*$$

(13)

where $p^*$ is the equilibrium price level, i.e. the price level which would be obtained at the actual money holdings if production and velocity were at their equilibrium levels. To determine $v^*$, a long-run money demand function (14) is estimated:

$$(m - p) = \beta y - v_0 + u$$

(14)

where $v_0$ is a constant. This is then solved with respect to $-(m - p - y) = v$ and hence for $v^*$. Finally $v^*$ is substituted again in the quantity equation (13) which yields an expression for $p^*$ in model (15) as an inverted money demand equation:

$$p^* = m + v_0 - \beta y^*$$

(15)

As demonstrated, inter alia, in Deutsche Bundesbank (1992a), in Issing and Toedter (1994) and in Toedter and Reimers (1994), this is exactly the Bundesbank’s way of deriving $p^*$ within the so-called P-Star concept of Hallman, Porter and Small (1991). Obviously model (14) is a

---

13 Ericsson’s N(0,1) test statistic clearly rejects that (Bbk) encompasses model (8) (6.3 [0-000]) but is significant at 5% in the opposite direction (1-99 [0-03]) and Cox’s test statistic rejects encompassing in either direction (see Doornik and Hendry, 1994, for details of the tests).
restricted version of model (1). However, imposing zero restrictions on inflation, interest rates and trend and re-estimating model (3), those restrictions are clearly rejected by $\chi^2(4) = 20.642$ [0.0004]. This contradicts the long-run specification by Toedter and Reimers (1994), who explain the downward trending behaviour of velocity solely by an income elasticity greater than unity. As shown in Figure 5, it is the (long-run) cointegration relationships (3) or (4) of those variables that explains the downward trend in velocity. Nevertheless, calculating $p^*$ still makes sense if $y$ could be treated as exogenous in the long run such that long-run forecasts for $y_t$ would not depend on forecasts for $(m - p)_t$, and if long-run variations in inflation and interest rates would be relatively small. However, following Hoover (1991) and Beyer (1993) the evidence that inflation is super-exogenous with respect to the parameters of interest of the money demand model seems to be in conflict with the assumption that prices change due to changes in money which is the key assumption of the P-Star approach. The question arises as through which transmission mechanism monetary policy kept inflation in Germany relatively low over the last 20 years. In order to answer this question it is necessary to have a more complete macroeconometric model in which more than just a money demand function is specified. A possible transmission mechanism which might be consistent with the statistical results found here might involve effects of monetary policy on prices via effects of interest rates on real goods demand and exchange rates and thence on prices via an

$\chi^2(4) = 20.642$ [0.0004]
aggregate supply relationship. Structural instability in the aggregate supply or goods demand relationships might be capable of masking a statistical causal link running from money to prices.

5. CONCLUSIONS

In this paper a money demand model for M3 in Germany was estimated. Applying an extended quarterly data set from 1975 to 1994 and hence including German unification in 1990, it was shown that the coefficients of an earlier pre-unification model remain empirically stable. Applying various tests for parameter instability, the null hypothesis of no structural break was not rejected although there is evidence that there are parameter shifts in the marginal processes of the super-exogenous explanatory variables inflation and long-term interest rates, respectively. It was therefore not rejected that the Lucas critique does not apply to the money demand model. The estimated model is based on SA data and conclusions drawn from it are robust to the use of NSA data. The model encompasses two recently published money demand models. The estimation of money demand as a single-equation model was justified because within a system the error-correction term of money demand does not enter the equations for the other explanatory variables. Furthermore, lags of real money do not Granger-cause inflation and long- and short-term interest rates. For further research, the empirical findings might suggest extending the analysis to estimating an econometric system to find more empirical evidence about the transmission channel of the Bundesbank’s monetary policy.

APPENDIX: DATA DEFINITIONS AND SOURCES

<table>
<thead>
<tr>
<th>Variable</th>
<th>Definition</th>
<th>Source</th>
</tr>
</thead>
<tbody>
<tr>
<td>$m$</td>
<td>Nominal M3 (SA); until 1990(4) for West Germany only</td>
<td>Deutsche Bundesbank</td>
</tr>
<tr>
<td>$m^u$</td>
<td>Nominal M3 (NSA); until 1990(1) for West Germany only</td>
<td>Datastream BDTU0800A1</td>
</tr>
<tr>
<td>$y$</td>
<td>Real GDP in 1991 prices (SA); until 1990(4) for West Germany only</td>
<td>Deutsche Bundesbank</td>
</tr>
<tr>
<td>$y^u$</td>
<td>Real GDP in 1991 prices (NSA); until 1991(4) for West Germany only</td>
<td>Statistisches Bundesamt</td>
</tr>
<tr>
<td>$yn$</td>
<td>Nominal GDP (SA); until 1990(4) for West Germany only</td>
<td>Datastream BDGDP..B</td>
</tr>
<tr>
<td>$p$</td>
<td>GDP(SA) deflator</td>
<td>(Implicit)</td>
</tr>
<tr>
<td>$p^u$</td>
<td>GDP(NSA) deflator</td>
<td>(Implicit)</td>
</tr>
<tr>
<td>$wT$</td>
<td>Total net financial wealth (domestic households and domestic enterprises, excluding housing); constructed and interpolated from annual flows; until 1990(4) for West Germany only</td>
<td>Deutsche Bundesbank</td>
</tr>
<tr>
<td>$wH$</td>
<td>Net financial wealth of domestic households; constructed and interpolated from annual flows; until 1990(4) for West Germany only</td>
<td>Deutsche Bundesbank</td>
</tr>
<tr>
<td>$RS$</td>
<td>Frankfurt three-month funds money market rate</td>
<td>Deutsche Bundesbank</td>
</tr>
<tr>
<td>$RL$</td>
<td>Yield on public debt securities outstanding</td>
<td>Deutsche Bundesbank</td>
</tr>
</tbody>
</table>
ACKNOWLEDGEMENTS

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