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Money, Interest Rate Spreads, and Economic Activity

By

Christine Sauer and Joachim Scheide

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I. Introduction

In recent years, the discussion about indicators of monetary policy has experienced a veritable renaissance, prompted by the breakdown of the basic money-income relationship in several industrial countries. The debate has been particularly lively in the United States, where the rapid pace of financial deregulation and innovation in the 1980s has advanced the process of financial disintermediation more than in any other industrial country. The resulting instability of the U.S. money demand function has led the Federal Reserve to deemphasize monetary aggregates and to search for a “better” indicator variable that is predictably linked to the final targets of output growth and price stability. One such variable that has attracted considerable attention is the term structure of interest rates, usually measured as the difference between long- and short-term nominal interest rates.

Given the long and variable lags between the implementation of monetary policy measures and their ultimate macroeconomic effects, two aspects are important when considering the qualities of a particular indicator variable. First, the variable should reflect the stance of monetary policy by signalling the direction and strength of policy actions. Second, the variable should act as an early indicator of future developments in the target variables to allow mid-course corrections if necessary.

As far as the term structure of interest rates is concerned, the policy indicator aspect has received relatively little attention in the

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recent U.S. literature.¹ There are, however, a number of European studies that consider this issue (e.g., Hesse and Roth 1992; Issing 1994; Ragnitz 1994). They generally argue against the term structure as an indicator variable because it not only reflects monetary policy but also factors such as inflationary and exchange rate expectations. While monetary policy can directly influence short-term interest rates, its influence on long-term interest rates (via inflationary expectations) is indirect at best. Furthermore, in a world of highly mobile capital, long-term interest rates are also influenced by international capital flows. Consequently, the term structure may not accurately reflect the stance of domestic monetary policy.

By contrast, the question whether the term structure is a reliable early indicator has received considerable attention. Numerous empirical studies for the United States and other industrial countries demonstrate that interest rate spreads are good predictors of both future economic growth² and future changes in the inflation rate (e.g., Mishkin 1990a, 1990b; Jorion and Mishkin 1991). Most studies adopt a bivariate model to compare the forecasting performance of the term structure to that of other indicator variables such as a stock-market index, a short-term interest rate, or different monetary aggregates. However, few researchers consider the question whether the term structure contains any information over and above what is already contained in other monetary variables. This issue is of considerable interest to policymakers in those countries where the money-income relationship has remained stable, implying that the empirical foundation for a successful policy of monetary targeting has not (yet) been undermined.

This paper contributes to the ongoing debate by considering a multivariate framework to assess the marginal information content of the term structure of interest rates with respect to economic activity. Methodological issues are discussed in Section II, whereas Section III presents the empirical evidence. Concluding remarks are given in Section IV.

II. Methodology

The question of information or predictive content is closely linked to the concept of Granger causality. In the presence of cointegration,

¹ Exceptions are Bernanke (1990) and Fuhrer (1993).

² See, for instance, Bernanke (1990), Estrella and Hardouvelis (1991), Harvey (1991), Friedman and Kuttner (1992), Hu (1993), Döpke and Gern (1993) and Langfeldt (1994).

causality tests must be based on a dynamic error-correction model (ECM) so that information about the long-run relationships between variables is taken into account (Granger 1988).³ In the bivariate case, the following “conditional model” for Δy is estimated:⁴

$$\Delta y_t = a_{10} + a_{11}(y_{t-1} - \beta x_{t-1}) + \sum_{i=1}^m b_{1i} \Delta y_{t-i} + \sum_{i=1}^n c_{1i} \Delta x_{t-i} + \varepsilon_{1t}, \quad (1)$$

where the term in parentheses is the lagged (stationary) error from the static cointegration regression $y_t = \beta x_t + u_t$. The hypothesis that x does not cause y must be rejected if the coefficient on the error-correction term a_{11} is significant, regardless of the joint significance of the c_{1i} -coefficients. Note that $a_{11} < 0$ implies that x and y are cointegrated.

The single-equation ECM for Δy can be estimated efficiently by least squares only if x is weakly exogenous with respect to the cointegration parameter (Engle et al. 1983). Tests for weak exogeneity are based on the following “marginal model” for Δx :

$$\Delta x_t = a_{20} + a_{21}(y_{t-1} - \beta x_{t-1}) + \sum_{i=1}^r b_{2i} \Delta x_{t-i} + \sum_{i=1}^s c_{2i} \Delta y_{t-i} + \varepsilon_{2t}, \quad (2)$$

where the null hypothesis of weak exogeneity imposes the restriction $a_{21} = 0$.⁵ This implies that x does not depend on the error-correction term calculated from the conditional model (1).

The methodology described here can be used to address the questions whether the term structure of interest rates contains information about economic activity over and above what is contained in other variables such as the readily observable monetary aggregates. To that end, we first consider a bivariate model to investigate the causal links between real output and real money.⁶ After adding the interest rate

³ Note that cointegration necessarily implies the existence of causality in at least one direction (Granger 1988, p. 203).

⁴ The error-correction specification is obtained by reparameterizing the rational lag model $B(L)y_t = C(L)x_t + u_t$, where $B(L) = 1 - B_1 L - \dots - B_{m+1} L^{m+1}$ and $C(L) = C_1 L + \dots + C_{n+1} L^{n+1}$ are polynomials in the lag operator. Consequently, the coefficients in (1) are linear combinations of those in the underlying rational lag model.

⁵ Reverse causality from y to x can be ruled out if $a_{21} = 0$ and $c_{2i} = 0$, that is, if x is strongly exogenous.

⁶ There are theoretical and statistical reasons for using money in real terms. Theoretical models indicate that *real* rather than *nominal* balances play a key role in the transmission of monetary policy impulses to the real sector. Many empirical studies indirectly capture this real balance effect by separately including nominal money growth and

spread to the list of explanatory variables, the tests for predictive content are repeated. Finally, to check the robustness of our findings, the same approach is applied to a trivariate model that includes the terms of trade.

III. Empirical Analysis

1. The Data

The analysis focuses on three large European countries – France, Germany, and Italy. The quarterly data run through 1994:2; the beginning of the sample period for each country is dictated by the availability of the various monetary aggregates (1978:1 for France, 1969:2 for Germany, and 1975:2 for Italy). In the German case, the data are for West Germany until June 1990 and for unified Germany thereafter.

The measure of economic activity is real domestic spending (real GDP minus net exports). Alternative definitions of money are the narrow aggregate M1 or a broader aggregate (M2 for Italy, M3 for France and Germany);⁷ in the case of Germany, central bank money (CBM) is used in addition. Real monetary aggregates are based on the deflator for domestic spending. The term structure of interest rates is defined as the difference between long- and short-term nominal interest rates. Short-term rates are measured by the 3-month money market rate (Germany), the 3-month interbank offered rate (France) or the rate on interbank sight deposits (Italy); the long-term rate is the yield on government or public/semi-public sector bonds. The terms of trade are calculated as the ratio of the export and import price deflators. Except for the interest rates, the data are seasonally adjusted at the source and expressed in terms of natural logarithms.⁸

inflation in the real output growth equation (e.g., Friedman and Kuttner 1992). This approach, however, is problematic since it mixes stationary and nonstationary variables, thus invalidating conventional testing procedures based on standard asymptotic theory (Sims et al. 1990). More specifically, we find that inflation and *nominal* money growth (except for Germany) are nonstationary whereas *real* money and real output are first-difference stationary for all three countries. (Results are available on request; see also Table 1.)

⁷ The raw money stock data are end-of-period monthly observations. The quarterly figures are computed as weighted averages from the monthly data; for example, the first quarter is 1/6 December + 1/3 January + 1/3 February + 1/6 March.

⁸ The data for France and Italy are from OECD *Main Economic Indicators*. The German data sources are Deutsche Bundesbank *Monatsberichte*, Statistisches Bundesamt *Wirtschaft und Statistik*, and DIW *Vierteljahreshefte zur Wirtschaftsforschung*.

2. Preliminary Tests

In the presence of cointegration, the dynamic relationship between output and various information variables can be represented by an error-correction model (Engle and Granger 1987). To determine the correct model specification for the causality analysis, it is thus necessary to test for cointegration. A prerequisite for cointegration, in turn, is that the variables in question are integrated of the same order. Unit root test results are presented in Table 1. For all three countries, the interest rate spread is found to be stationary in levels [I(0)], whereas the other variables are stationary after first-differencing [I(1)]. Consequently, the term structure of interest rates does not enter the long-run cointegration relationship(s) between output, money, and the terms of trade.

Next, the Engle-Granger two-step procedure is used to perform preliminary cointegration tests on the I(1)-variables. The evidence in Table 2 suggests that the hypothesis of no cointegration must be rejected for all model specifications and money definitions in Germany, but not in France and Italy. However, the residual-based cointegration test has been shown to possess reduced power because it imposes the "common factor restriction" of identical short- and long-run elasticities (Kremers et al. 1992). A more powerful test is obtained by estimating the error-correction model directly and checking the significance of the error-correction term.

3. Error-Correction Estimates and Granger Causality Tests

As a first step, we analyze the relationships between economic activity (Y) and various real monetary aggregates (RM) by estimating the following single-equation error-correction model (Model 1):

$$\Delta Y_t = a_0 + a_1 Y_{t-1} + a_2 RM_{t-1} + \sum_{i=1}^p b_i \Delta Y_{t-i} + \sum_{i=1}^p c_i \Delta RM_{t-i} + \varepsilon_t, \quad (3)$$

where the "optimal" lag length p is chosen on the basis of the Akaike Information Criterion (AIC). The German equations also include a dummy variable ($\text{dummy} = 1$ for 1990:3 and 1990:4) to account for the effects of the reunification and some statistical changes in the monetary aggregates (Deutsche Bundesbank 1991). The information content of real money is assessed by testing the hypothesis $c_1 = \dots = c_p = 0$ by means of a simple F-test.

Table 1 – Unit Root Tests

| Log of variable ^a | Level | | | Difference | | | Result |
|-----------------------------------|----------|------------------|--------------------|------------|------------------|-----------------|--------|
| | <i>p</i> | ADF ^b | PP ^b | <i>p</i> | ADF ^b | PP ^b | |
| <i>France</i> (1978:1–1994:2) | | | | | | | |
| Real domestic spending | 3 | -2.28 | -1.50 | 0 | -7.13** | -7.39** | I(1) |
| Real M1 | 1 | -1.23 | -1.37 | 0 | -5.92** | -6.10** | I(1) |
| Real M3 | 9 | -1.69 | -1.59 | 0 | -3.96** | -3.99** | I(1) |
| Terms of trade | 1 | -2.41 | -1.98 | 0 | -5.85** | -5.98** | I(1) |
| Short-term interest rate | 1 | -0.62 | -0.82 | 0 | -6.30** | -6.30** | I(1) |
| Long-term interest rate | 1 | -0.55 | -0.41 | 0 | -4.79** | -4.76** | I(1) |
| Interest rate spread | 0 | -2.11* | -2.47* | | NA | NA | I(0) |
| <i>Italy</i> (1975:2–1994:2) | | | | | | | |
| Real domestic spending | 2 | -2.40 | -1.98 | 0 | -6.29** | -6.50** | I(1) |
| Real M1 | 1 | -1.88 | -2.12 | 0 | -5.51** | -5.62** | I(1) |
| Real M2 | 1 | -1.49 | -1.52 | 0 | -5.25** | -5.38** | I(1) |
| Terms of trade | 0 | -2.24 | -2.34 | 0 | -8.13** | -8.23** | I(1) |
| Short-term interest rate | 0 | -0.56 | -0.61 | 0 | -7.47** | -7.50** | I(1) |
| Long-term interest rate | 1 | -0.47 | -0.35 | 0 | -5.30** | -5.37** | I(1) |
| Interest rate spread | 1 | -3.68** | -3.79** | | NA | NA | I(0) |
| <i>Germany</i> (1969:2–1994:2) | | | | | | | |
| Real domestic spending | 3 | -1.96 | -1.55 | 0 | -8.93** | -9.18** | I(1) |
| Real M1 | 2 | -1.59 | -1.10 | 1 | -4.65** | -7.34** | I(1) |
| Real M3 | 1 | -1.60 | -1.43 | 0 | -8.30** | -8.33** | I(1) |
| Real central bank money | 0 | -0.58 | -1.04 | 1 | -4.93** | -8.69** | I(1) |
| Terms of trade | 2 | -1.90 | -1.95 | 1 | -5.35** | -8.10** | I(1) |
| Short-term interest rate | 1 | -1.26 | -2.14 ⁺ | 0 | -6.08** | -6.01** | I(1) |
| Long-term interest rate | 3 | -0.59 | -1.07 | 2 | -4.32** | -6.93** | I(1) |
| Interest rate spread | 1 | -3.41** | -3.18** | | NA | NA | I(0) |

^a Except for interest rates and interest rate spreads. – ^b ADF = augmented Dickey-Fuller t-test (with *p* lags to ensure that the residuals are free of first and fourth order autocorrelation). PP = Phillips-Perron t-test (based on 4 covariance lags). For levels, the ADF-regression includes a constant and a linear time trend; for growth rates, the trend is omitted. In the case of interest rates and interest rate spreads, the regression includes neither a constant nor a time trend. **, *, and ⁺ indicate rejection of the unit root hypothesis at the 1, 5, and 10 percent significance levels, respectively, based on the critical values from Fuller (1976, Table 8.5.2, p. 373).

Table 2 – *Engle-Granger Cointegration Tests*

| Dep. | Indep. | France | | Italy | | Germany | |
|-------------|------------------|----------|---------------------|----------|---------------------|----------|---------------------|
| | | <i>p</i> | t-test ^a | <i>p</i> | t-test ^a | <i>p</i> | t-test ^a |
| <i>Y</i> | <i>RM1</i> | 0 | -1.25 | 1 | -1.87 | 3 | -3.67* |
| <i>Y</i> | <i>RM3</i> | 0 | -2.86 | 1 | -2.15 | 4 | -3.16 ⁺ |
| <i>Y</i> | <i>RCBM</i> | | | | | 3 | -3.64* |
| <i>RM1</i> | <i>Y</i> | 1 | -1.33 | 1 | -1.58 | 3 | -3.51* |
| <i>RM3</i> | <i>Y</i> | 0 | -2.64 | 1 | -1.39 | 4 | -3.14 ⁺ |
| <i>RCBM</i> | <i>Y</i> | | | | | 3 | -3.45* |
| <i>Y</i> | <i>TOT, RM1</i> | 0 | -1.36 | 1 | -2.15 | 3 | -3.59 ⁺ |
| <i>Y</i> | <i>TOT, RM3</i> | 0 | -2.78 | 1 | -2.71 | 4 | -3.47 ⁺ |
| <i>Y</i> | <i>TOT, RCBM</i> | | | | | 3 | -3.73* |
| <i>RM1</i> | <i>Y, TOT</i> | 0 | -1.32 | 1 | -1.75 | 3 | -3.43 ⁺ |
| <i>RM3</i> | <i>Y, TOT</i> | 1 | -2.69 | 1 | -1.58 | 4 | -3.41 ⁺ |
| <i>RCBM</i> | <i>Y, TOT</i> | | | | | 3 | -3.53 ⁺ |

^aThe augmented Dickey-Fuller test is based on $\Delta \hat{u}_t = \gamma \hat{u}_{t-1} + \sum_{i=1}^p \delta_i \Delta \hat{u}_{t-i} + e_t$, where \hat{u}_t is the estimated residual from the long-run model $y_t = \alpha + \beta x_t + u_t$. The lag length p is chosen so that the residuals e_t are free of first- and fourth-order autocorrelation. **, *, and ⁺ indicate rejection of the hypothesis of no cointegration at the 1, 5, and 10 percent significance levels, respectively, based on the critical values from Engle and Yoo (1987, Tables 2 and 3, pp. 157–158).

To determine whether the term structure contains any additional information, we add the difference between long- and short-term interest rates (ID) and estimate the augmented specification (Model 2), using the same lag structure as before:

$$\Delta Y_t = a_0 + a_1 Y_{t-1} + a_2 RM_{t-1} + \sum_{i=1}^p b_i \Delta Y_{t-i} + \sum_{i=1}^p c_i \Delta RM_{t-i} + \sum_{i=1}^p e_i ID_{t-i} + \varepsilon_t. \quad (4)$$

Again, F-tests are performed to assess the information content of real money ($c_1 = \dots = c_p = 0$) and the term structure ($e_1 = \dots = e_p = 0$) with respect to economic activity. A significant F-test for the e_t -coefficients implies that the yield curve cannot be excluded from the model. Similarly, if the inclusion of ID results in a lower standard error of the estimate (SEE) and/or a lower final prediction error (FPE), the term structure provides information over and above what is already contained in real money.

The error-correction specification allows a direct test for cointegration based on the significance of the error-correction term (Kremers et al. 1992); if $a_1 < 0$, the hypothesis of no cointegration must be rejected. The cointegration parameter can be calculated as $\beta_1 \equiv -a_2/a_1$.⁹ Note that β_1 can be estimated efficiently with the single-equation ECM for ΔY as long as the other variables are weakly exogenous. The exogeneity tests are based on the marginal models consistent with the conditional models (3) and (4).¹⁰

Table 3 reports the estimation and hypothesis test results for the bivariate and trivariate ECMs. In France, the inclusion of the interest rate spread does not improve the fit of the ΔY -equation, and the ID-coefficients are jointly insignificant (Model 2). Consequently, the term structure does not contain information over and above what is already captured by money. Based on the significance of the error-correction term, the hypothesis of no cointegration must be rejected for the narrow monetary aggregate. This implies the existence of a causal relationship between economic activity and real M1 despite the fact that the c_i -coefficients are jointly insignificant. No such link is detected for the broad aggregate M3. The diagnostic tests confirm that the estimated residuals are white noise for all specifications. However, the hypothesis of weak exogeneity must be rejected for real M3, implying that estimates based on the single-equation ECM are not efficient.

For Italy, we also find that the interest rate spread does not contain any additional information regarding future economic activity (Model 2). The hypothesis of no cointegration must be rejected for all specifications. Since the coefficients on ΔRM are jointly insignificant, the causal link between economic activity and real money is limited to the common trend that the two variables share. The diagnostic check reveals that the residuals from Model 1 may not be free of fourth-order ARCH effects; in Model 2, the hypothesis of no first-order serial correlation must be rejected at the 10 percent level. For all four specifications, the regressors in the output equation are found to be weakly exogenous.

⁹ The corresponding t-statistic is obtained from the instrumental variable estimation of the Bewley transformed equation.

¹⁰ For example, the marginal model for ΔRM consistent with (3) is $\Delta RM_t = \alpha_0 + \alpha_1 EC_{t-1} + \sum_{i=1}^p \gamma_i \Delta RM_{t-i} + \sum_{i=1}^p \delta_i \Delta Y_{t-i} + \eta_t$, where EC is the error-correction term calculated from the conditional model for ΔY . Weak exogeneity imposes the restriction $\alpha_1 = 0$.

Table 3 – Error-Correction Estimates and Granger Causality Tests for France, Italy, and Germany

| | France | | | | Italy | | | | Germany | | | | | |
|-------------------------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|---------------------|--------------------|--------------------|---------------------|--------------------|--------------------------------|
| | Model 1 | | Model 2 | | Model 1 | | Model 2 | | Model 1 | | | Model 2 | | |
| | RM1 | RM3 | RM1 | RM3 | RM1 | RM2 | RM1 | RM2 | RM1 | RM3 | RCBM | RM1 | RM3 | RCBM |
| No. of lags | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 3 | 3 | 3 | 3 | 3 | 3 |
| EC-term | -0.049 (2.643)* | -0.064 (1.192) | -0.054 (2.099)* | -0.061 (1.129) | -0.041 (2.288)* | -0.031 (2.316)* | -0.038 (2.047)* | -0.028 (2.025)* | -0.176 (4.151)** | -0.081 (2.442)* | -0.092 (2.555)* | -0.170 (3.609)** | -0.035 (0.956) | -0.070 (1.958) ⁺ |
| Cointegration coefficient β_1 | 1.942 (4.703)** | 0.554 (4.302)** | 1.844 (3.772)** | 0.654 (3.744)** | 1.005 (2.538)* | 0.979 (1.665) | 1.006 (2.316)* | 0.997 (1.538) | 0.639 (24.334)** | 0.619 (9.768)** | 0.638 (9.836)** | 0.656 (22.525)** | 0.679 (4.651)** | 0.727 (8.555)** |
| F-test: $c_i=0$ | 1.24 | 1.92 | 1.28 | 1.01 | 0.90 | 0.24 | 0.59 | 0.07 | 3.23* | 1.03 | 2.38 ⁺ | 3.31* | 1.11 | 2.76* |
| F-test: $e_i=0$ | - | - | 0.08 | 0.64 | - | - | 0.30 | 0.64 | - | - | - | 1.06 | 2.90* | 3.47* |
| <i>Diagnostics</i> | | | | | | | | | | | | | | |
| R ² | 0.1638 | 0.0645 | 0.1649 | 0.0747 | 0.1757 | 0.1562 | 0.1793 | 0.1639 | 0.6082 | 0.4955 | 0.5335 | 0.6226 | 0.5429 | 0.5850 |
| SEE | 0.0083 | 0.0088 | 0.0084 | 0.0088 | 0.0113 | 0.0115 | 0.0114 | 0.0115 | 0.0112 | 0.0127 | 0.0122 | 0.0112 | 0.0123 | 0.0117 |
| FPE $\times 10e^{-4}$ | 0.7710 | 0.8626 | 0.7947 | 0.8805 | 1.4102 | 1.4436 | 1.4422 | 1.4693 | 1.4090 | 1.8144 | 1.6776 | 1.4455 | 1.7507 | 1.5895 |
| LM(1)-test | 0.38 | 0.50 | 0.22 | 1.38 | 2.69 | 2.29 | 3.38 ⁺ | 3.10 ⁺ | 1.62 | 2.75 ⁺ | 1.50 | 0.98 | 1.32 | 1.14 |
| LM(4)-test | 2.09 | 3.86 | 2.01 | 3.02 | 4.95 | 5.01 | 6.96 | 6.81 | 7.26 | 7.97 ⁺ | 6.28 | 8.29 ⁺ | 9.51* | 8.47 ⁺ |
| ARCH(1)-test | 1.03 | 1.45 | 1.13 | 1.43 | 0.45 | 0.47 | 1.00 | 1.06 | 0.05 | 0.31 | 0.84 | 0.41 | 0.46 | 0.46 |
| ARCH(4)-test | 4.06 | 3.62 | 4.11 | 3.00 | 8.62 ⁺ | 10.03* | 7.18 | 8.07 ⁺ | 1.41 | 0.81 | 0.90 | 2.18 | 0.68 | 1.83 |
| Normality test | 2.05 | 0.85 | 1.91 | 1.05 | 1.60 | 1.58 | 1.29 | 1.15 | 0.67 | 0.94 | 0.98 | 0.53 | 0.36 | 0.54 |
| <i>Weak exogeneity tests</i> | | | | | | | | | | | | | | |
| F-test: RM | 0.72 | 4.11* | 2.52 | 6.70* | 0.18 | 0.37 | 0.45 | 0.40 | 1.84 | 0.24 | 0.00 | 0.96 | 1.38 | 0.41 |
| F-test: ID | - | - | 0.14 | 0.75 | - | - | 1.39 | 1.16 | - | - | - | 0.68 | 1.47 | 0.27 |
| Joint LR-test | - | - | 2.78 | 7.93* | - | - | 2.81 | 2.53 | - | - | - | 1.56 | 3.23 | 0.92 |

Note: Figures in parentheses are absolute t-statistics. SEE is the standard error of the estimate, FPE is the final prediction error. The LM-statistics test for first- and fourth-order residual correlation; the ARCH-statistics test whether the squared residuals follow a first- or fourth-order autoregressive process; the normality of the residuals is tested by means of the Jarque-Bera statistic. Tests for weak exogeneity are based on the corresponding marginal model(s), using an F-test for single equations and a likelihood ratio (LR) test for a system of equations. Significance at the 1, 5, and 10 percent levels is indicated by **, *, and ⁺, respectively.

Surprisingly, the German results are most supportive of the notion that the term structure of interest rates contains useful information about future economic activity once real money is taken into account. This is true for the broad aggregate M3 and the central bank money stock (CBM), as indicated by the significant F-statistics on the e_i -coefficients and the considerably lower SEE and FPE in Model 2. The null hypothesis of no cointegration must be rejected in five out of six cases, so that causality exists in at least one direction. In the case of M1 and CBM, the jointly significant c_i -coefficients indicate that causality also runs from changes in real money to economic activity. The estimated residuals are normally distributed and free of ARCH effects; for some specifications, however, the hypothesis of no serial correlation must be rejected at the 10 percent significance level. Since the regressors in the output equations are found to be weakly exogenous, the single-equation approach is appropriate in the German case.

One potential shortcoming of the models tested so far is that they do not explicitly incorporate real disturbances. Terms of trade shocks, in particular, may contribute significantly to short-run output fluctuations in large open economies.¹¹ Scheide (1989), for example, finds that the West German business cycle is influenced by both monetary and terms of trade fluctuations.

Preliminary tests indicate that the terms of trade (TOT) contain a unit root (see Table 1), thus satisfying a prerequisite for cointegration with real domestic spending and real money. If these variables are cointegrated, the long-run equilibrium relationship must be taken into account when evaluating the information content of different monetary variables.¹² Incorporating the terms of trade, we first estimate the following error-correction model (Model 3):

¹¹ One might, of course, extend the information set even further to avoid a potential omitted variable problem. Real exports, for example, have been found to Granger-cause productivity in Germany (Marin 1992). Our results for Germany, France, and Italy (not reported, but available on request) indicate that no such causal link exists between real exports and real domestic spending. Furthermore, we find no evidence of cointegration among real domestic spending, real exports, and the terms of trade. For these reasons, real exports are excluded from the subsequent analysis. In general, however, the information content of one variable for forecasting another can be evaluated only with respect to the information set included in the model.

¹² The results in Table 2 suggest that the hypothesis of no cointegration in the trivariate system must be rejected for Germany, but not for France and Italy. However, the Engle-Granger cointegration test is not as powerful as the direct test based on the error-correction specification (Kremers et al. 1992).

$$\begin{aligned} \Delta Y_t = & a_0 + a_1 Y_{t-1} + a_2 RM_{t-1} + a_3 TOT_{t-1} + \sum_{i=1}^p b_i \Delta Y_{t-i} \\ & + \sum_{i=1}^p c_i \Delta RM_{t-i} + \sum_{i=1}^p d_i \Delta TOT_{t-i} + \varepsilon_t. \end{aligned} \quad (5)$$

Next, we estimate the augmented specification that includes the interest rate spread as an additional regressor (Model 4):

$$\begin{aligned} \Delta Y_t = & a_0 + a_1 Y_{t-1} + a_2 RM_{t-1} + a_3 TOT_{t-1} \\ & + \sum_{i=1}^p b_i \Delta Y_{t-i} + \sum_{i=1}^p c_i \Delta RM_{t-i} + \sum_{i=1}^p d_i \Delta TOT_{t-i} \\ & + \sum_{i=1}^p e_i ID_{t-i} + \varepsilon_t. \end{aligned} \quad (6)$$

As before, the “optimal” lag length p is determined on the basis of the Akaike Information Criterion (AIC). Predictive content is assessed by means of a simple F-test and the fit of the model (SEE, FPE). The evidence for the three countries is reported in Table 4.

Overall, the results are remarkably robust with respect to this change in specification. Germany remains the only country for which the term structure of interest rates contains information about future economic activity that is not already captured by real money or the terms of trade (Model 4). Based on the significance of the error-correction term, the hypothesis of no cointegration must be rejected in ten out of fourteen cases. This implies that a causal link exists between economic activity, real money, and the terms of trade regardless of the significance of the short-run coefficients. The diagnostic tests generally confirm that the estimated residuals are white noise; the exogeneity tests indicate that the single-equation specification is appropriate in all but two cases.

IV. Concluding Remarks

The question whether the term structure of interest rates is a reliable early indicator of economic activity has been discussed extensively in the recent theoretical and empirical literature. This paper contributes to the ongoing debate by considering a multivariate framework to assess the information content of different monetary variables. The analysis is based on the error-correction specification of short-run economic fluctuations to incorporate the causal links that exist between real and monetary variables in the long run.

Table 4 – Error-Correction Estimates and Granger Causality Tests for France, Italy, and Germany

| | France | | | | Italy | | | | Germany | | | | | |
|-------------------------------------|---------------------|-------------------|--------------------|-------------------|--------------------|--------------------|--------------------|--------------------|---------------------|---------------------|--------------------|---------------------|--------------------|--------------------|
| | Model 3 | | Model 4 | | Model 3 | | Model 4 | | Model 3 | | | Model 4 | | |
| | RM1 | RM3 | RM1 | RM3 | RM1 | RM2 | RM1 | RM2 | RM1 | RM3 | RCBM | RM1 | RM3 | RCBM |
| No. of lags | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 1 | 3 | 3 | 3 | 3 | 3 | 3 |
| EC-term | -0.055 (2.893)** | -0.028 (0.450) | -0.065 (2.436)* | -0.030 (0.481) | -0.045 (2.182)* | -0.040 (2.322)* | -0.043 (2.033)* | -0.039 (2.262)* | -0.166 (3.878)** | -0.098 (2.785)** | -0.086 (2.384)* | -0.140 (2.885)** | -0.045 (1.185) | -0.057 (1.652) |
| Cointegration coefficient β_1 | 1.958 (3.616)** | 0.127 (0.097) | 1.814 (3.606)** | 0.307 (0.318) | 0.724 (1.228) | 0.379 (0.503) | 0.613 (0.935) | 0.170 (0.211) | 0.641 (23.344)** | 0.655 (13.166)** | 0.657 (9.856)** | 0.676 (17.975)** | 0.746 (5.647)** | 0.842 (6.088)** |
| Cointegration coefficient β_2 | 0.083 (0.170) | 1.121 (0.356) | -0.101 (0.244) | 0.919 (0.356) | 0.474 (0.656) | 0.905 (1.104) | 0.664 (0.820) | 1.225 (1.416) | 0.073 (0.599) | 0.514 (2.134)* | 0.293 (1.114) | 0.153 (0.934) | 1.087 (1.245) | 0.801 (1.370) |
| F-test: $c_i=0$ | 0.39 | 1.00 | 0.44 | 0.68 | 0.28 | 0.00 | 0.06 | 0.18 | 3.37* | 1.33 | 2.55 ⁺ | 4.00** | 2.21 ⁺ | 4.52** |
| F-test: $d_i=0$ | 2.17 | 1.19 | 2.38 | 1.07 | 0.03 | 0.06 | 0.05 | 0.09 | 1.65 | 1.54 | 1.95 | 1.98 | 2.11 | 2.46 ⁺ |
| F-test: $e_i=0$ | - | - | 0.31 | 0.16 | - | - | 0.54 | 1.23 | - | - | - | 1.73 | 3.75* | 5.42** |
| <i>Diagnostics</i> | | | | | | | | | | | | | | |
| R ² | 0.1946 | 0.1012 | 0.1990 | 0.1037 | 0.1803 | 0.1669 | 0.1868 | 0.1819 | 0.6329 | 0.5541 | 0.5747 | 0.6553 | 0.6091 | 0.6466 |
| SEE | 0.0083 | 0.0088 | 0.0084 | 0.0088 | 0.0115 | 0.0116 | 0.0115 | 0.0116 | 0.0111 | 0.0122 | 0.0119 | 0.0109 | 0.0116 | 0.0111 |
| FPE $\times 10e^{-4}$ | 0.7911 | 0.8828 | 0.8122 | 0.9088 | 1.4799 | 1.5040 | 1.5083 | 1.5174 | 1.4358 | 1.7441 | 1.6634 | 1.4372 | 1.6295 | 1.4732 |
| LM(1)-test | 1.17 | 1.41 | 1.33 | 0.96 | 2.67 | 1.77 | 3.40 ⁺ | 2.59 | 1.12 | 0.89 | 0.68 | 0.18 | 0.38 | 1.25 |
| LM(4)-test | 2.56 | 4.39 | 2.73 | 3.46 | 6.80 | 3.87 | 9.16 ⁺ | 6.24 | 6.73 | 5.28 | 4.41 | 9.10 ⁺ | 8.23 ⁺ | 9.07 ⁺ |
| ARCH(1)-test | 0.14 | 0.48 | 0.14 | 0.51 | 0.16 | 0.03 | 0.64 | 0.56 | 0.07 | 0.83 | 0.88 | 0.34 | 0.93 | 0.75 |
| ARCH(4)-test | 1.73 | 1.87 | 1.47 | 1.80 | 9.81* | 9.67* | 8.31 ⁺ | 7.37 | 1.45 | 1.83 | 1.66 | 1.34 | 1.51 | 2.27 |
| Normality test | 1.18 | 1.18 | 1.06 | 1.12 | 1.87 | 1.88 | 1.53 | 1.32 | 1.84 | 3.59 | 2.65 | 4.32 | 6.19* | 4.27 |
| <i>Weak exogeneity tests</i> | | | | | | | | | | | | | | |
| F-test: <i>RM</i> | 0.22 | 0.46 | 1.55 | 2.40 | 0.25 | 0.31 | 0.44 | 0.99 | 1.73 | 1.40 | 0.01 | 2.30 | 2.03 | 2.40 |
| F-test: <i>TOT</i> | 0.55 | 5.42* | 0.52 | 5.62* | 1.19 | 0.63 | 1.52 | 0.85 | 0.01 | 0.07 | 0.00 | 0.06 | 1.02 | 0.23 |
| F-test: <i>ID</i> | - | - | 0.00 | 0.01 | - | - | 2.65 | 2.02 | - | - | - | 0.01 | 0.02 | 0.05 |
| Joint LR-test | 0.67 | 5.70 ⁺ | 1.87 | 7.36 ⁺ | 2.14 | 1.59 | 5.10 | 4.27 | 2.23 | 2.28 | 0.01 | 3.29 | 5.74 | 3.94 |

Note: See Table 3.

The evidence indicates that the information content of the term structure, over and above what is already captured by real monetary aggregates, differs across countries. In France and Italy, the difference between long- and short-term interest rates does not improve forecasts of real domestic spending; it does, however, in the case of Germany. By contrast, previous studies that adopt a bivariate setting generally find the term structure to be a good predictor of economic activity across countries.

What are the practical implications of the findings presented here? Market participants and policymakers in Germany might be well advised to include the term structure of interest rates in their business cycle forecast models. As for the conduct of monetary policy, the (limited) early indicator properties of the term structure should not be misconstrued to recommend the abolishment of monetary targeting in favor of interest rate targeting. The arguments against such a switch in intermediate targets are both theoretical and empirical. First, a given interest rate spread may be associated with high or low inflation. Second, a stable relationship between money and prices continues to exist, for example, in Germany (e.g., Deutsche Bundesbank 1992; Krämer and Scheide 1994). The money stock provides an anchor for the price level, a fact that can be used to implement a policy compatible with the desired inflation rate. The term structure of interest rates does not possess this property.

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Abstract: Money, Interest Rate Spreads, and Economic Activity. – Numerous empirical studies for industrial countries have shown that the term structure of interest rates is a good indicator for future output growth. This paper analyzes whether the interest rate spread contains any additional predictive power if the model includes the money stock. A multivariate error-correction framework is applied to three European economies – France, Germany, and Italy. Granger causality tests are performed for various monetary aggregates and the term structure. The evidence concerning the marginal information content is mixed: For France and Italy, the term structure does not improve the results of the basic model whereas it is significant for Germany. JEL No. E32, E44.

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Zusammenfassung: Geldmenge, Zinsdifferenz und wirtschaftliche Aktivität. – Zahlreiche empirische Studien für Industriestaaten zeigen, daß die Differenz zwischen lang- und kurzfristigen Zinsen ein guter Frühindikator für die gesamtwirtschaftliche Produktion ist. Der vorliegende Aufsatz untersucht für drei westeuropäische Länder (Deutschland, Frankreich und Italien), ob die Zinsdifferenz die Prognosequalität des Modells verbessert, das eine Geldmengengröße enthält. Dazu werden multivariate Fehlerkorrekturmodelle geschätzt und Tests auf Granger-Kausalität durchgeführt. Die Evidenz hinsichtlich des marginalen Informationsgehalts der Zinsdifferenz ist unterschiedlich: Für Frankreich und Italien wird die Prognosegüte des Basismodells durch die Hinzunahme dieser Variablen nicht verbessert, während für Deutschland ein signifikanter Einfluß vorliegt.
