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MONEY, INTEREST RATE SPREADS, AND ECONOMIC ACTIVITY

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Abstract

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 addresses the question whether the interest rate spread contains any *additional* predictive power if information on the money stock is already included in the model. A multivariate error-correction framework is applied to three large European economies – France, Germany, and Italy. The importance of various (real) monetary aggregates and the term structure is investigated with Granger causality tests. The models also include the terms of trade as an indicator of real disturbances. The evidence concerning the marginal information content of the interest rate spread is mixed. For France and Italy, the variable does not improve the results of the basic model whereas it plays a significant role in the case of Germany.
We conclude that policy makers and market participants should check carefully whether the term structure can improve business cycle forecasts. Regardless of its indicator qualities, however, the variable should not be viewed as a possible intermediate target for monetary policy. The term structure does not provide an anchor for the price level and, thus, is not an alternative to monetary targeting.

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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 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 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

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

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 growth [e.g., Bernanke (1990), Estrella and Hardouvelis (1990), Harvey (1991), Friedman and Kuttner (1992), Hu (1993), Döpke and Gern (1993), Langfeldt (1994)] 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 policy variables. This issue is of considerable interest to policy makers 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, 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. A stationary variable x is said to Granger–cause another stationary variable y if past values of x contain information about current y over and above the information contained in past values of y alone. Since most macroeconomic time series are known to contain a unit root, x and y are usually rendered stationary by first–differencing before causality tests are performed on Δx and Δy. These standard tests, however, are misspecified if x and y share a common trend, that is, if they are cointegrated [Granger (1988), p. 204]. In this case, the temporal relationship between the two variables can be represented by an error–correction model (ECM) that incorporates past deviations from the common trend as an additional channel through which x can influence current Δy [Engle and Granger (1987)]. Note that cointegration necessarily implies the existence of causality in at least one direction [Granger (1988), p. 203].

More formally, the modified Granger causality tests are based on the following dynamic error–correction model for Δy ("conditional model"):[2]

$$(1) \qquad \Delta y_t = a_{10} + a_{11}(y_{t-1} - \beta x_{t-1}) + \sum_{i=1}^{p} b_{1i}\Delta y_{t-i} + \sum_{i=1}^{q} c_{1i}\Delta x_{t-i} + \epsilon_{1t},$$

where the term in parentheses is the lagged (stationary) error from the static cointegration regression $y_t = \beta x_t + u_t$. In this framework, the hypothesis that x

[2]For the bivariate case, the error–correction specification is obtained by reparameterizing the rational lag model $B(L)y_t = C(L)x_t + u_t$, where $B(L)=1-B_1L-\ldots-B_{p+1}L^{p+1}$ and $C(L)=C_1L+\ldots+C_{q+1}L^{q+1}$ are polynomials in the lag operator. Consequently, the coefficients in equation (1) are linear combinations of those in the underlying rational lag model.

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 β [Engle, Hendry, and Richard (1983)]. Tests for weak exogeneity are based on the following error–correction model for Δx ("marginal model"):

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

where the null hypothesis of weak exogeneity imposes the restriction $a_{21}=0$.[3] This implies that x does not depend on the error–correction term calculated from the conventional model (1).

The methodology described here for the bivariate case can be used to address the question whether the term structure of interest rates contains information about economic activity over and above what is contained in other indicators of monetary policy such as different monetary aggregates. To that end, we first consider a bivariate model to analyze the causal links between money and output. Next, the interest rate spread is added to the list of explanatory variables and the tests for predictive content are repeated. Alternatively, the spread variable is added to a trivariate model that includes the terms of trade to capture any real disturbances.[4]

[3]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.

[4]Several authors find that the terms of trade have a significant impact on output [e.g., Scheide (1989)].

III. EMPIRICAL ANALYSIS

A. The Data

The analysis focuses on three large European countries —France, Germany, and Italy— that are characterized by relatively stable money–income relationships.[5] 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);[6] in addition, central bank money (CBM) is used in the case of Germany. 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, which capture real disturbances to the economy, 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

[5]Reviewing almost 400 multicountry studies, Fase (1994) concludes that Germany has the most stable money demand among the large EU countries, followed by France and Italy.

[6]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.

of natural logarithms.[7]

B. Preliminary Tests

In the presence of cointegration, the dynamic relationship between output and various indicators of monetary policy can be represented by an error–correction model (ECM). To determine the correct model specification for the Granger 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. The 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 all other variables are stationary after first–differencing [I(1)].[8] Consequently, the term structure of interest rates does not enter the long–run cointegration relationship(s) between output, money, and the terms of trade.

(Table 1 about here.)

Next, the Engle–Granger two–step procedure is used to perform preliminary cointegration tests on the I(1)–variables.[9] Test results are reported in Table 2 below. In the case of Germany, the hypothesis of no cointegration must be rejected

[7]The data for France and Italy are from OECD Main Economic Indicators. The German data sources are Deutsche Bundesbank, Statistisches Bundesamt, and DIW.

[8]The evidence for Germany indicates that the short– and long–term interest rates may be I(0).

[9]This residual–based cointegration test has been shown to possess reduced power because it imposes the "commom factor restriction" of identical short– and long–run elasticities. A more powerful test for cointegration is obtained by estimating the error–correction model in one step and checking the significance of the error–correction term. See Kremers et al. (1992).

for all model specifications and money definitions. In France, the null can only be rejected for the broad monetary aggregate M3. For Italy, there is some evidence of cointegration in the bivariate but not in the trivariate model. In summary, the preliminary cointegration tests indicate that the link between output, money, and the terms of trade should be modeled by an error–correction specification.

(Table 2 about here.)

C. Error–Correction Estimates and Granger Causality Tests

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

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

where the "optimal" lag length p is chosen on the basis of the Akaike Information Criterion (AIC). Note that the German equations also include a dummy variable (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 money is assessed by testing the hypothesis $c_1=...=c_p=0$ by means of a simple F–test.

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:

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

Again, F–tests are performed to assess the information content of real money ($c_1=...=c_p=0$) and the term structure ($e_1=...=e_p=0$) with respect to economic activity. A significant F–test for the e_i–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), then 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, which does not impose the "common factor restriction" implicit in the Engle–Granger two–step procedure [Kremers et al. (1992)]. The hypothesis of no cointegration must be rejected if the coefficient on the error–correction term a_1 is significantly negative. The cointegration parameter can be calculated as $\beta_1 \equiv -a_2/a_1$.[10] 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).[11]

[10]The corresponding t–statistic is obtained from the instrumental variable estimation of the Bewley transformed equation.

[11]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 $\alpha_1=0$.

(Tables 3a–3c about here.)

Tables 3a, 3b, and 3c report 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; for Model 2, the hypothesis of no first–order serial correlation is rejected only at the 5% level. For all four specifications, the regressors in the output equation are found to be weakly exogenous.

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, the hypothesis of no serial correlation can only be rejected at higher significance levels. 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 explicitely incorporate real disturbances. Terms of trade shocks, in particular, may contribute significantly to short–run output fluctuations in large open economies. 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 analyzing the information content of different indicators of monetary policy.[12] Adopting the same approach as before, we incorporate the terms of trade by first estimating the following single–equation error–correction model (Model 3):

[12]The results in Table 2 suggest that the hypothesis of no cointegration in the trivariate system must be rejected for France and Germany, but not for 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)].

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

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

$$(6)\quad \Delta Y_t = a_0 + a_1 Y_{t-1} + a_2 RM_{t-1} + a_3 TOT_{t-1} + \sum_{i=1}^{p} b_i \Delta Y_{t-i} + \sum_{i=1}^{p} c_i \Delta RM_{t-i} + \sum_{i=1}^{p} d_i \Delta TOT_{t-i} + \sum_{i=1}^{p} e_i ID_{t-i} + \epsilon_t,$$

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 estimates for the three countries are reported in Tables 4a, 4b, and 4c.

(Tables 4a–4c about here.)

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 indicators of monetary policy. The analysis is based on error–correction models of short–run economic fluctuations to incorporate the causal links that exist between real and monetary variables in the long run.

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 policy makers 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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Table 1.
Unit Root Tests

Log of Variable[a]	Level p	Level ADF[b]	Level PP[b]	Difference p	Difference ADF[b]	Difference PP[b]	Result
France (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	-2.65	-2.03	0	-6.26**	-6.28**	I(1)
Long-Term Interest Rate	1	-2.72	-2.07	0	-4.74**	-4.75**	I(1)
Interest Rate Spread	0	-3.31+	-3.59*	5	-5.46**	-7.86**	I(0)
Italy (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	1	-2.72	-2.30	0	-7.42**	-7.49**	I(1)
Long-Term Interest Rate	1	-2.52	-2.07	0	-5.26**	-5.36**	I(1)
Interest Rate Spread	1	-4.30**	-3.84*	1	-7.35**	-8.19**	I(0)
Germany (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	- .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	-3.36+	-2.88	0	-6.04**	-6.00**	I(1)
Long-Term Interest Rate	3	-3.16+	-2.66	2	-4.30**	-6.94**	I(1)
Interest Rate Spread	1	-3.92*	-3.20+	0	-6.42**	-6.35**	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). Regressions include a constant and (for levels) a linear time trend. **, *, and + indicate rejection of the unit root hypothesis at the 1%, 5%, and 10% significance level, respectively, based on the critical values from Fuller (1976, Table 8.5.2, p. 373).

Table 2.
Engle-Granger Cointegration Tests

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

[a]The augmented Dickey-Fuller test (with p lags to ensure that the residuals are free of first and fourth order autocorrelation) is performed on the estimated residual $\hat{u}_t$ from the long-run model $y_t = \alpha + \beta x_t + u_t$. **, *, and + indicate rejection of the hypothesis of no cointegration at the 1%, 5%, and 10% significance level, respectively, based on the critical values from Engle and Yoo (1987, Tables 2 and 3, pp. 157-58).

Table 3a.
Error-Correction Estimates & Granger-Causality Tests: France

	Model 1		Model 2	
	RM1	RM3	RM1	RM3
No. of lags	1	1	1	1
EC-term	-.049 (2.643)*	-.064 (1.192)	-.054 (2.099)*	-.061 (1.129)
Cointegration coefficient β_1	1.942 (4.703)**	.554 (4.302)**	1.844 (3.772)**	.654 (3.744)**
F-test: c_i=0	1.24	1.92	1.28	1.01
F-test: e_i=0	-	-	.08	.64
Diagnostics				
R^2	.1638	.0645	.1649	.0747
SEE	.0083	.0088	.0084	.0088
FPE*10e-4	.7710	.8626	.7947	.8805
LM(1)-test	.38	.50	.22	1.38
LM(4)-test	2.09	3.86	2.01	3.02
ARCH(1)-test	1.03	1.45	1.13	1.43
ARCH(4)-test	4.06	3.62	4.11	3.00
Normality test	2.05	.85	1.91	1.05
Weak exogeneity tests				
F-test: RM	.72	4.11*	2.52	6.70*
F-test: ID	-	-	.14	.75
Joint LR-test	-	-	2.78	7.93*

Notes: Figures in parentheses are absolute t-statistics. SEE is the standard error of the estimate, FPE 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 a single equation and a likelihood ratio (LR) test for a system of equations. Significance at the 1%, 5%, and 10% levels is indicated by **, *, and +, respectively.

Table 3b.
Error-Correction Estimates & Granger-Causality Tests: Italy

	Model 1		Model 2	
	RM1	RM2	RM1	RM2
No. of lags	1	1	1	1
EC-term	-.041 (2.288)*	-.031 (2.316)*	-.038 (2.047)*	-.028 (2.025)*
Cointegration coefficient β_1	1.005 (2.538)*	.979 (1.665)	1.006 (2.316)*	.997 (1.538)
F-test: c_i=0	.90	.24	.59	.07
F-test: e_i=0	-	-	.30	.64
Diagnostics				
R^2	.1757	.1562	.1793	.1639
SEE	.0113	.0115	.0114	.0115
FPE*10e-4	1.4102	1.4436	1.4422	1.4693
LM(1)-test	2.69	2.29	3.38+	3.10+
LM(4)-test	4.95	5.01	6.96	6.81
ARCH(1)-test	.45	.47	1.00	1.06
ARCH(4)-test	8.62+	10.03*	7.18	8.07+
Normality test	1.60	1.58	1.29	1.15
Weak exogeneity tests				
F-test: RM	.18	.37	.45	.40
F-test: ID	-	-	1.39	1.16
Joint LR-test	-	-	2.81	2.53

Notes: See table 3a.

Table 3c.
Error-Correction Estimates & Granger-Causality Tests: Germany

	Model 1			Model 2		
	RM1	RM3	RCBM	RM1	RM3	RCBM
No. of lags	3	3	3	3	3	3
EC-term	-.176 (4.151)**	-.081 (2.442)*	-.092 (2.555)*	-.170 (3.609)**	-.035 (.956)	-.070 (1.958)+
Cointegr. coeff. β_1	.639 (24.334)**	.619 (9.768)**	.638 (9.836)**	.656 (22.525)**	.679 (4.651)**	.727 (8.555)**
F-test: c_i=0	3.23*	1.03	2.38+	3.31*	1.11	2.76*
F-test: e_i=0	-	-	-	1.06	2.90*	3.47*
Diagnostics						
R^2	.6082	.4955	.5335	.6226	.5429	.5850
SEE	.0112	.0127	.0122	.0112	.0123	.0117
FPE*10e-4	1.4090	1.8144	1.6776	1.4455	1.7507	1.5895
LM(1)-test	1.62	2.75+	1.50	.98	1.32	1.14
LM(4)-test	7.26	7.97+	6.28	8.29+	9.51*	8.47+
ARCH(1)-test	.05	.31	.84	.41	.46	.46
ARCH(4)-test	1.41	.81	.90	2.18	.68	1.83
Normality test	.67	.94	.98	.53	.36	.54
Weak exogeneity tests						
F-test: RM	1.84	.24	.00	.96	1.38	.41
F-test: ID	-	-	-	.68	1.47	.27
Joint LR-test	-	-	-	1.56	3.23	.92

Notes: See table 3a.

Table 4a.
Error-Correction Estimates & Granger-Causality Tests: France

	Model 3		Model 4	
	RM1	RM3	RM1	RM3
No. of lags	1	1	1	1
EC-term	-.055 (2.893)**	-.028 (.450)	-.065 (2.436)*	-.030 (.481)
Cointegration coefficient β_1	1.958 (3.616)**	.127 (.097)	1.814 (3.606)**	.307 (.318)
Cointegration coefficient β_2	-.083 (.170)	1.121 (.356)	-.101 (.244)	.919 (.356)
F-test: c_i=0	.39	1.00	.44	.68
F-test: d_i=0	2.17	1.19	2.38	1.07
F-test: e_i=0	-	-	.31	.16
Diagnostics				
R^2	.1946	.1012	.1990	.1037
SEE	.0083	.0088	.0084	.0088
FPE*10e-4	.7911	.8828	.8122	.9088
LM(1)-test	1.17	1.41	1.33	.96
LM(4)-test	2.56	4.39	2.73	3.46
ARCH(1)-test	.14	.48	.14	.51
ARCH(4)-test	1.73	1.87	1.47	1.80
Normality test	1.18	1.18	1.06	1.12
Weak exogeneity tests				
F-test: RM	.22	.46	1.55	2.40
F-test: TOT	.55	5.42*	.52	5.62*
F-test: ID	-	-	.00	.01
Joint LR-test	.67	5.70+	1.87	7.36+

Notes: See table 3a.

Table 4b.
Error-Correction Estimates & Granger-Causality Tests: Italy

	Model 3		Model 4	
	RM1	RM2	RM1	RM2
No. of lags	1	1	1	1
EC-term	-.045 (2.182)*	-.040 (2.322)*	-.043 (2.033)*	-.039 (2.262)*
Cointegration coefficient β_1	.724 (1.228)	.379 (.503)	.613 (.935)	.170 (.211)
Cointegration coefficient β_2	.474 (.656)	.905 (1.104)	.664 (.820)	1.225 (1.416)
F-test: c_i=0	.28	.00	.06	.18
F-test: d_i=0	.03	.06	.05	.09
F-test: e_i=0	-	-	.54	1.23
Diagnostics				
R^2	.1803	.1669	.1868	.1819
SEE	.0115	.0116	.0115	.0116
FPE*10e-4	1.4799	1.5040	1.5083	1.5174
LM(1)-test	2.67	1.77	3.40+	2.59
LM(4)-test	6.80	3.87	9.16+	6.24
ARCH(1)-test	.16	.03	.64	.56
ARCH(4)-test	9.81*	9.67*	8.31+	7.37
Normality test	1.87	1.88	1.53	1.32
Weak exogeneity tests				
F-test: RM	.25	.31	.44	.99
F-test: TOT	1.19	.63	1.52	.85
F-test: ID	-	-	2.65	2.02
Joint LR-test	2.14	1.59	5.10	4.27

Notes: See table 3a.

Table 4c.
Error-Correction Estimates & Granger-Causality Tests: Germany

	Model 3			Model 4		
	RM1	RM3	RCBM	RM1	RM3	RCBM
No. of lags	3	3	3	3	3	3
EC-term	-.166 (3.878)**	-.098 (2.785)**	-.086 (2.384)*	-.140 (2.885)**	-.045 (1.185)	-.057 (1.652)
Cointegr. coeff. β_1	.641 (23.344)**	.655 (13.166)**	.657 (9.856)**	.676 (17.975)**	.746 (5.647)**	.842 (6.088)**
Cointegr. coeff. β_2	.073 (.599)	.514 (2.134)*	.293 (1.114)	.153 (.934)	1.087 (1.245)	.801 (1.370)
F-test: c_i=0	3.37*	1.33	2.55+	4.00**	2.21+	4.52**
F-test: d_i=0	1.65	1.54	1.95	1.98	2.11	2.46+
F-test: e_i=0	-	-	-	1.73	3.75*	5.42**
Diagnostics						
R^2	.6329	.5541	.5747	.6553	.6091	.6466
SEE	.0111	.0122	.0119	.0109	.0116	.0111
FPE*10e-4	1.4358	1.7441	1.6634	1.4372	1.6295	1.4732
LM(1)-test	1.12	.89	.68	.18	.38	1.25
LM(4)-test	6.73	5.28	4.41	9.10+	8.23+	9.07+
ARCH(1)-test	.07	.83	.88	.34	.93	.75
ARCH(4)-test	1.45	1.83	1.66	1.34	1.51	2.27
Normality test	1.84	3.59	2.65	4.32	6.19*	4.27
Weak exogeneity tests						
F-test: RM	1.73	1.40	.01	2.30	2.03	2.40
F-test: TOT	.01	.07	.00	.06	1.02	.23
F-test: ID	-	-	-	.01	.02	.05
Joint LR-test	2.23	2.28	.01	3.29	5.74	3.94

Notes: See table 3a.

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