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ABSTRACT

What Drives the European Central Bank's Interest-Rate Changes?*

We show that the ECB's interest rate changes during 1999-2010 have been mainly driven by changes in economic activity in the Euro area. Changes in actual or expected future HICP inflation play a minor, if any, role.

JEL Classification: E43 and E58
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1. Introduction

Since Taylor (1993) showed how a simple linear combination of inflation and GDP deviations from trend to some extent mimicked the path of the U.S. Federal Funds rate in the period 1987Q1-92Q4, a voluminous literature has assessed monetary policy conduct through what is now known as *the Taylor rule*. From a practical and normative perspective, policy deliberations in many central banks include the Taylor rule as a yardstick for good policymaking.[1] Also, theoretical research and model-based forecasting often use Taylor-type rules as a default representation of monetary policymaking, or as a potential approximation to optimal policy; see Galí (2008) for a recent textbook exposition.

An important characteristic of the rule is that the nominal interest rate should be raised more than proportionally when inflation raises—this is often labelled an *active* Taylor rule. This response is viewed as important, as it is conducive for inflation stability. For example, it precludes sun-spot-driven equilibria in forward-looking economies. The empirical literature has thus often focused on whether the rule is active or not (e.g., Clarida *et al.,* 2000).

The objective of this note is to assess whether the Taylor rule has had any relevance for monetary policy in the Euro area since its inception in 1999. Our empirical evidence shows that it has not. Despite the fact that inflation stability is the primary objective of the European Central Bank, the ECB has not responded systematically to inflation. Instead, the main determinant of interest-rate changes, within a class of simple response functions, is economic activity, e.g., unemployment changes in the Euro area. We argue that this does not need to be an indication of problematic monetary policy conduct—neither in terms of stability issues, nor in terms of optimality considerations.

2. Theory, estimation strategy and data

In our main formulation of a testable behavioral expression for the ECB, we follow the literature and allow for the inclusion of expected future variables to capture potential forward-looking aspects in monetary policy; see, e.g., Gerdesmeier and Roffia (2006), Gerlach (2007) and Gerlach and Lewis (2010) for reviews of the empirical literature on ECB behavior. A forward-looking Taylor-type setting of the target value for the nominal interest

[1] Asso *et al.* (2010) present an intriguing account of the history of the Taylor rule, and show how it is mentioned several times in the transcripts of FOMC meetings. Also, in policy reports of the inflation-targeting Central Bank of Norway, the Taylor rule is mentioned as input to assessing the "appropriate interest rate"; see Norges Bank (2010, pp. 22–25).

rate i_t^*, is usually modelled as:

$$i_t^* = \alpha + \beta \left(\mathrm{E}_t \pi_{t+k} - \pi^* \right) + \gamma \mathrm{E}_t \left(y_{t+h} - y_{t+h}^n \right) + \boldsymbol{\delta} \mathbf{x}_t', \tag{1}$$

where π_t in the inflation rate, π^* is the goal value for inflation, y_t is output, y_t^n is the natural rate of output, $\mathbf{x}_t$ is a vector of other variables that may influence interest-rate setting, E_t is the rational expectations operator conditional on information available in t, and $(\alpha, \pi^*, \beta, \gamma, \boldsymbol{\delta})$ is the vector of parameters to be estimated. The specification allows for both $k > 0$ and $h > 0$, i.e., that the ECB may respond to expectations about future inflation and the output gap. As is standard is the literature, autocorrelation in observed nominal interest rates motivates estimation of a partial adjustment model:

$$i_t = \sum_{j=1}^{m} \rho_j i_{t-j} + \left(1 - \sum_{j=1}^{m} \rho_j \right) i_t^* + \widetilde{\varepsilon}_t, \qquad m > 0, \tag{2}$$

where $\sum_{j=1}^{m} \rho_j$ is interpreted as the degree of "interest rate smoothing," and $\widetilde{\varepsilon}_t$ is an i.i.d. "policy shock."

Initial examinations of (2) gave results that were very sensitive to specification. One reason could be that data are nonstationary (or near integrated), which would render parameter estimates spurious. Indeed, it is not possible to reject nonstationarity of most of our variables; cf. the Augmented Dickey-Fuller tests in our Supplementary Appendix 4. This property of data has been acknowledged in the empirical literature, but most researchers nevertheless assess versions of (2) arguing that the unit-root tests have low power in short samples. While we are sympathetic towards this argument, the difficulty of obtaining robust results led us to estimate (2) in first differences, where data according to the ADF tests are likely to be stationary.[2] Thereby, we are unable to identify α and π^*, but as our interest is monetary policy responses at the business cycle frequency, lack of identification of these long-run parameters is less important. In the case of $m = 2$ (the highest value we found significant), (2) then becomes

$$\Delta i_t = \rho_1 \Delta i_{t-1} + \rho_2 \Delta i_{t-2} + (1 - \rho_1 - \rho_2) \Delta i_t^* + \Delta \widetilde{\varepsilon}_t$$

[2] Christensen and Nielsen (2009) use co-integrating techniques in estimations of a Taylor-type rule on US data. Österholm (2005) performs co-integration analysis on standard Taylor rules for US and other countries. We have not pursued co-integration analysis here, as most theory does not envisage a Taylor rule as a long-run representation of interest rate determination.

(Δ is the differencing operator), and thus by (1):

$$\Delta i_t = \rho_1 \Delta i_{t-1} + \rho_2 \Delta i_{t-2} + (1-\rho_1-\rho_2)\left[\beta \Delta \mathrm{E}_t \pi_{t+k} + \gamma \Delta \mathrm{E}_t \left(y_{t+h} - y^n_{t+h}\right) + \boldsymbol{\delta} \Delta \mathbf{x}'_t\right] + \Delta \widetilde{\varepsilon}_t. \tag{3}$$

We estimate (3) by General Methods of Moments (Hansen, 1982) as the right-hand side variables are endogenous and/or unknown in period t.[3] Replacing expected values by actual values, the relation becomes

$$\Delta i_t = \rho_1 \Delta i_{t-1} + \rho_2 \Delta i_{t-2} + (1-\rho_1-\rho_2)\left[\beta \Delta \pi_{t+k} + \gamma \Delta y_{t+h} + \boldsymbol{\delta} \Delta \mathbf{x}'_t\right] + \varepsilon_t, \tag{4}$$

where $\varepsilon_t \equiv (1-\rho_1-\rho_2)\,\beta\Delta\left(\mathrm{E}_t\pi_{t+k} - \pi_{t+k}\right) + (1-\rho_1-\rho_2)\,\gamma\Delta\left(\mathrm{E}_t y_{t+h} - y_{t+h} - \mathrm{E}_t y^n_{t+h}\right) + \Delta\widetilde{\varepsilon}_t$. (Note that the difference specification avoids the problem of the unobserved natural rate of output.) The parameter vector $\boldsymbol{\theta} \equiv (\rho_1, \rho_2, \beta, \gamma, \boldsymbol{\delta})$ is then identified using the orthogonality, or, moment conditions, $\mathrm{E}_t\left[\varepsilon_t \mathbf{z}_t\right] = 0$, where $\mathbf{z}_t$ is a vector of instrument variables known in period t.[4] The parameter vector is exactly identified when the number of instrument variables match the number of parameters, but we use more instruments, which allows for an evaluation of the validity of the specification and instruments by the Hansen J-test for overidentifying restrictions. We use relatively small sets of instruments, as it is well known that too many instruments makes it virtually impossible to reject the validity of moment conditions. As instruments we use lags of the variables in the equation with lag lengths suggested by the univariate pattern of autocorrelation. An optimal weighting of moment conditions is adopted by an iterative procedure, and we use Heteroskedasticity and Autocorrelation Consistent covariance estimators adopting the Andrews-Monahan method with $AR(1)$ pre whitening.[5]

We use monthly data covering the period from 1999m1 to 2010m1. As ECB's nominal interest rate, we follow most empirical literature and use the Euro Overnight Index Average (EONIA); cf. Gerdesmeier and Roffia (2006) and Gerlach and Lewis (2010). As inflation we use the measure that is ECB's self-proclaimed goal variable, the rate of change in the Harmonized Index of Consumer Prices (HICP). As output we use Euro-16 GDP, which,

[3]For estimation, we apply Michael T. Cliff's MATLAB routines (Cliff, 2003), which is kindly made available at http://www.feweb.vu.nl/econometriclinks/mcliffprogs.html. Our programs and data are available upon request.

[4]All estimations involve a constant term (as do the vectors of instruments), which is ignored here and in the ensuing tables as it is always insignificant.

[5]We have chosen this method over the Newey-West procedure with a Bartlett kernel, where the bandwidth is to be chosen. First, Andrews and Monahan (1992) show that other HAC estimators bias t-statistics upwards. Second, the significance of some variables was sensitive to the choice of bandwidth. Hence, we use Andrews and Monahan's estimator—which automatically determines the bandwidth—so as to avoid rejecting H_0 hypotheses too often, and to avoid taking a stand on bandwidth choice.

however, is only available on a quarterly basis. For the estimations, we therefore extract monthly values by a cubic spline transformation. We subsequently consider specifications with unemployment, u_t, as activity variable—this data is available for Euro-16 at monthly frequency.[6] The Supplementary Appendix B contains full details on all data sources.

3. Results

Table 1 presents the parameter estimates for (4) at various $k \geq h$. This choice of leads are motivated by the conventional wisdom that monetary policy changes first impact on output and then later on inflation. Apart from Hansen's J-test, we also report the adjusted R^2 and the Ljung-Box Q-test for autocorrelation (up to, and including, 18 lags). These diagnostics are not pertinent to GMM estimation (and note that a negative R^2 is possible), but we include them to give some information about the fit and behavior of the residuals of the estimated relations. The information set reflects that HICP inflation is only known with a month's lag and unemployment with two months' lag.

The estimations based on observable output reveal one consistent pattern, namely that inflation never enters significantly (and only in the instance of a six month lead is the estimated parameter value in conformity with an active Taylor rule). Output enters significantly in some instances, but the point estimates varies substantially across specifications (also, the specifications suffer from either autocorrelation of errors or high J-statistics). The estimations based on expectations of current output, deliver strongly significant estimates of γ, and the point estimates are quite similar across specifications. Inflation at any horizon enters insignificantly. The specifications including expected future output all perform rather poorly. In the case where both output and inflation are six-month ahead expectations, we obtain a significant estimate of the impact of inflation. The point estimate is very high, 3.23, but the specification suffers from strong autocorrelation in residuals, and a lack of explanatory power. Also, in the other specifications where inflation enters as six-month ahead expectations, point estimates are very different (–0.41 and 1.58, respectively). Finally, interest rate smoothing is virtually absent; some specifications portray a significant impact of interest-rate changes two months before, but these mostly occurs in specifications where output and inflation are insignificant. All in all, the most robust finding is the significant impact of current output changes on current interest-rate changes, as well as the lack of any systematic effect of HICP inflation.

[6] Recent theory also emphasizes that unemployment is a relevant variable for monetary policy analysis; see Blanchard and Galí (2010).

Table 1: Changes in EONIA explained by leads and lags of HICP-inflation changes and GDP growth

ρ_1	ρ_2	$\beta_{k=-1}$	$\beta_{k=0}$	$\beta_{k=6}$	$\beta_{k=12}$	$\gamma_{h=-2}$	$\gamma_{h=0}$	$\gamma_{h=6}$	$\gamma_{h=12}$	$\overline{R}^2$	J	$Q_{(18)}$
0.13	0.17	0.05	—	—	—	$1.09^{\ddagger}$	—	—	—	0.43	—	18.67
(0.13)	(0.12)	(0.08)				(0.30)						(0.10)
0.12	$0.29^{\dagger}$	—	0.82	—	—	$1.06^{\dagger}$	—	—	—	0.11	10.16	25.85
(0.12)	(0.14)		(0.64)			(0.50)					(0.07)	(0.10)
0.09	$0.40^{\ddagger}$	—	—	1.58	—	$0.77^{\dagger}$	—	—	—	–0.77	3.31	43.89
(0.14)	(0.13)			(0.96)		(0.39)					(0.65)	(0.00)
–0.14	0.09	—	—	—	0.25	$1.91^{\ddagger}$	—	—	—	0.19	7.55	26.12
(0.14)	(0.11)				(0.18)	(0.23)					(0.18)	(0.10)
0.07	$0.27^{\ddagger}$	—	0.18	—	—	—	$1.28^{\ddagger}$	—	—	0.52	4.27	19.45
(0.10)	(0.08)		(0.21)				(0.21)				(0.51)	(0.36)
0.04	0.24	—	—	–0.41	—	—	$1.62^{\ddagger}$	—	—	0.32	4.46	30.73
(0.11)	(0.18)			(0.45)			(0.43)				(0.49)	(0.03)
0.01	$0.30^{\dagger}$	—	—	—	0.25	—	$1.69^{\ddagger}$	—	—	0.38	13.90	15.93
(0.10)	(0.12)				(0.21)		(0.33)				(0.02)	(0.60)
$0.20^{\dagger}$	$0.53^{\ddagger}$	—	—	$3.23^{\dagger}$	—	—	—	3.58	—	–1.66	1.95	63.02
(0.09)	(0.07)			(1.35)				(2.30)			(0.86)	(0.00)
0.22	$0.40^{\ddagger}$	—	—	—	–1.40	—	—	4.75	—	–0.92	5.20	32.75
(0.13)	(0.14)				(0.98)			(2.62)			(0.39)	(0.02)
0.17	$0.40^{\ddagger}$	—	—	—	0.36	—	—	—	0.57	0.15	9.19	29.30
(0.11)	(0.10)				(0.44)				(1.55)		(0.10)	(0.05)

Notes: GMM estimates of (4) at various $k \geq h$. (First line, however, reports OLS regression.) Standard errors in parentheses. Sample period: 1999m1–2010m1

The instrument set is $\mathbf{z}_t^y = \{1, \Delta i_{t-1}, \Delta i_{t-2}, \Delta i_{t-3}, \Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta y_{t-2}, \Delta y_{t-3}, \Delta y_{t-4}, \Delta y_{t-5}\}$. We adopt an iteratively updated weighting matrix, and Heteroskedasticity and Autocorrelation Consistent estimators using the Andrews and Monahan (1992) method with $AR(1)$ pre whitening. J is the Hansen test for the validity of overidentifying restrictions (with the associated P-value for rejecting validity in parenthesis). $Q_{(18)}$ is the Ljung-Box Q-test for autocorrelation of residuals up to 18 lags (with the associated P-value for rejecting no autocorrelation in parenthesis).

†Significant at the 5% level. ‡Significant at the 1% level.

Table 2: Changes in EONIA explained by leads and lags of HICP-inflation and unemployment changes

ρ_1	ρ_2	$\beta_{k=-1}$	$\beta_{k=0}$	$\beta_{k=6}$	$\beta_{k=12}$	$\gamma_{h=-2}$	$\gamma_{h=0}$	$\gamma_{h=6}$	$\gamma_{h=12}$	$\overline{R}^2$	J	$Q_{(18)}$
0.28†	0.28†	0.12	—	—	—	–0.65	—	—	—	0.36	—	19.10
(0.13)	(0.12)	(0.16)				(0.44)						(0.38)
0.16†	0.39‡	—	0.84	—	—	0.15	—	—	—	0.29	5.00	20.16
(0.12)	(0.11)		(0.74)			(0.75)					(0.42)	(0.32)
0.17	0.21	—	—	–0.21	—	–1.09†	—	—	—	0.25	6.92	23.07
(0.12)	(0.14)			(0.22)		(0.51)					(0.23)	(0.18)
0.19	0.26‡	—	—	—	0.20	–1.46‡	—	—	—	0.30	9.41	14.37
(0.13)	(0.09)				(0.19)	(0.47)					(0.09)	(0.70)
0.18	0.19	—	0.03	—	—	—	–1.27‡	—	—	0.44	2.85	22.76
(0.11)	(0.13)		(0.41)				(0.45)				(0.72)	(0.20)
0.14	0.20	—	—	0.02	—	—	–1.59‡	—	—	0.46	3.34	21.63
(0.11)	(0.14)			(0.18)			(0.38)				(0.65)	(0.25)
0.14	0.23	—	—	—	–0.01	—	–2.12‡	—	—	0.44	1.73	16.93
(0.11)	(0.13)				(0.19)		(0.36)				(0.88)	(0.53)
0.02	0.27†	—	—	–0.49†	—	—	—	–2.05‡	—	–0.09	3.66	39.09
(0.13)	(0.11)			(0.23)				(0.65)			(0.60)	(0.00)
0.17	0.32†	—	—	—	–0.40	—	—	–2.76‡	—	0.10	3.14	17.55
(0.12)	(0.15)				(0.31)			(0.78)			(0.68)	(0.48)
0.14	0.51‡	—	—	—	0.32	—	—	—	–0.21	0.23	16.14	16.44
(0.14)	(0.07)				(0.38)				(1.46)		(0.01)	(0.56)

Notes: GMM estimates of (4) at various $k \geq h$—with unemployment u_t replacing y_t. (First line, however, reports OLS regression.) Standard errors in parentheses. Sample period: 1999m1–2010m1

The instrument set is $\mathbf{z}_t^u = \{1, \Delta i_{t-1}, \Delta i_{t-2}, \Delta i_{t-3}, \Delta \pi_{t-1}, \Delta \pi_{t-2}, \Delta u_{t-2}, \Delta u_{t-3}, \Delta u_{t-4}, \Delta u_{t-5}\}$. We adopt an iteratively updated weighting matrix, and Heteroskedasticity and Autocorrelation Consistent estimators using the Andrews and Monahan (1992) method with $AR(1)$ pre whitening. J is the Hansen test for the validity of overidentifying restrictions (with the associated P-value for rejecting validity in parenthesis). $Q_{(18)}$ is the Ljung-Box Q-test for autocorrelation of residuals up to 18 lags (with the associated P-value for rejecting no autocorrelation in parenthesis).

†Significant at the 5% level. ‡Significant at the 1% level.

Since the output measure is derived from quarterly data, we make the same estimations in Table 2, except that output is replaced by unemployment for which we have monthly data. The results qualitatively corroborate those presented in Table 1. Current unemployment has a strongly significant impact on interest-rate setting, and the point estimates are comparable to those obtained with output as the activity variable. (HICP inflation only enters significantly in one specification; with a negative sign.) The results seem more robust compared to those based on output when it comes to the validity of the moment conditions. No matter at what horizon inflation enters, the J-tests show that one cannot reject the validity of the moment conditions when unemployment enters contemporaneously. Quantitatively, the results suggest that an increase in Euro-area unemployment by 1 percentage point is met by a decrease in the short interest rate by over 100 basis points. Again, consistent evidence on interest rate smoothing is hard to find—two-month lagged interest-rate changes have significant impact in some instances, but mostly when unemployment enters at an insignificant lag or lead.

To further examine interest-rate setting, we run the regression with expected current unemployment and HICP inflation and add (first differences of) other macroeconomic variables one at the time. In all cases, variables enter contemporaneously, and for those that are endogenous or unobservable at t we add their lagged values to the instrument set (the number of lags determined from univariate autoregressive properties). In each case, we also present the results for the case where we eliminate the insignificant HICP inflation. Table 3 reports the results, where column δ contains the estimated coefficient to the variable under consideration. Note that in the case without any additional variables, the specification is virtually unchanged when HICP inflation is removed from the equation.

Of some interest is the estimate of δ when $M3$ growth is added. It is *negative*, and significantly so when inflation is not included. This does not support that the ECB takes money growth into account when evaluating inflationary pressures, which contrasts with its official reliance of so-called "monetary analysis" (cf. Papademos and Stark, 2010).[7] The price of oil enters significantly and with the expected positive sign, but only in the case where inflation is absent (the magnitude is very modest though; a 50% increase in the oil price leads to a 20 basis point nominal interest-rate increase). The ECB's economic sentiment index enters significantly with the expected sign—inclusion of this activity measure, however, do not rule out an independent response to unemployment movements. Movements in exchange rates, Dollar/Euro rate or the effective real Euro

[7] To be fair, this analysis emphasizes the medium-term evolution of money aggregates, and not their business cycle component as we consider here.

Table 3: Adding other explanatory variables

	ρ_1	ρ_2	β	γ	δ	$\overline{R}^2$	J	$Q_{(18)}$
	0.18	0.19	0.03	–1.27‡	—	0.45	2.85	22.76
None	(0.11)	(0.13)	(0.41)	(0.45)			(0.72)	(0.20)
	0.20	0.20	—	–1.38‡	—	0.45	1.83	22.87
	(0.12)	(0.14)		(0.34)			(0.77)	(0.20)
	0.08	0.25	0.41	–1.42†	–0.22	0.29	4.94	23.88
$M3$ growth	(0.15)	(0.14)	(0.29)	(0.67)	(0.16)		(0.67)	(0.16)
	0.24	0.13	—	–2.24‡	–0.31†	0.24	4.12	24.32
	(0.15)	(0.15)		(0.49)	(0.14)		(0.66)	(0.15)
	0.17	0.17	0.24	–1.24‡	0.003	0.47	4.12	27.56
Oil price	(0.09)	(0.11)	(0.30)	(0.35)	(0.004)		(0.66)	(0.07)
	0.15	0.22†	—	–1.34‡	0.005†	0.48	3.24	27.51
	(0.10)	(0.10)		(0.25)	(0.002)		(0.66)	(0.07)
	0.16	0.15	0.00	–1.55‡	0.036‡	0.47	7.70	25.26
Economic senti-	(0.10)	(0.13)	(0.37)	(0.45)	(0.012)		(0.17)	(0.12)
ment index	0.11	0.18	—	–1.57‡	0.037‡	0.47	6.27	24.28
	(0.10)	(0.12)		(0.24)	(0.012)		(0.18)	(0.14)
	0.14	0.20	0.11	–1.30‡	–0.009	0.45	2.19	22.74
Dollar/Euro	(0.19)	(0.11)	(0.36)	(0.49)	(0.029)		(0.82)	(0.20)
exchange rate	0.19	0.17	—	–1.52‡	0.008	0.45	1.38	22.97
	(0.13)	(0.13)		(0.42)	(0.037)		(0.85)	(0.19)
	0.14	0.20	–0.16	–1.24‡	–0.087	0.19	1.03	19.55
Real effective	(0.13)	(0.16)	(0.42)	(0.53)	(0.083)		(0.96)	(0.36)
exchange rate	0.11	0.22	—	–1.15‡	–0.076	0.30	1.18	18.33
	(0.12)	(0.16)		(0.41)	(0.079)		(0.88)	(0.36)
	0.19	0.15	–0.18	–1.46‡	–0.03	0.36	4.77	23.75
Long Euro-	(0.12)	(0.14)	(0.28)	(0.37)	(0.31)		(0.57)	(0.16)
bond rate	0.20	0.22	—	–1.29‡	0.32	0.41	3.74	27.94
	(0.13)	(0.16)		(0.39)	(0.46)		(0.59)	(0.06)
	0.19†	0.21†	0.20	–1.19†	–0.03	0.46	3.08	23.06
Lehman Brothers	(0.08)	(0.10)	(0.38)	(0.51)	(0.07)		(0.69)	(0.19)
Dummy 1	0.19†	0.26‡	—	–1.15‡	–0.10	0.43	1.76	20.60
	(0.09)	(0.09)		(0.41)	(0.13)		(0.78)	(0.30)
	0.20†	0.32†	0.18	–0.97	–1.55	0.51	2.38	28.89
Lehman Brothers	(0.09)	(0.13)	(0.48)	(0.55)	(0.91)		(0.79)	(0.05)
Dummy 2	0.22†	0.32†	—	–1.13‡	–1.87‡	0.51	2.17	29.35
	(0.10)	(0.13)		(0.43)	(0.71)		(0.71)	(0.04)
	0.03	0.32‡	0.32	–1.42‡	—	0.17	4.61	23.20
None —	(0.08)	(0.12)	(0.31)	(0.47)			(0.46)	(0.18)
sample ends in	0.02	0.29†	—	–1.49‡	—	0.22	4.03	25.81
June 2008	(0.08)	(0.13)		(0.40)			(0.40)	(0.10)

Notes: See notes to Table 2 for estimation method. For further details, see Appendix A. †Significant at the 5% level. ‡Significant at the 1% level.

rate, are not found to have any impact on interest-rate setting. The same applies for the long (10 year) Euro bond rate.

As our sample includes data from one of the most turbulent economic periods in recent decades, the financial crisis that erupted in the Fall of 2008, it is of concern to assess whether our main results are dependent on this event. Indeed many of the larger swings in the nominal interest rate occur during the crisis. This issue has recently been addressed by Gerlach and Lewis (2010) who estimate a Taylor-rule model allowing for endogenous regime shifts. They use data in levels and find a regime shift around September 2008. This coincides with the collapse of the Lehman Brothers. We therefore introduce a dummy capturing this shift. "Dummy 1" is results for the inclusion of a dummy that takes on zero up until September 2008, one half in September 2008, and one afterwards (interestingly, this crude dummy closely mimics the estimated indicator for regime shift in Gerlach and Lewis, 2010). The dummy variable is insignificant. As we conduct estimations in first differences, we also consider the dummy in differences—see the results for "Dummy 2". Here, the dummy is significant when inflation is not part of the equation (and has the expected negative sign), but unemployment is still a significant impetus for interest-rate changes (note that this specification suffers from strong autocorrelation in errors). To further examine the potential impact of the financial crisis, we reestimate the basic equation for a briefer sample ending in June 2008. As seen, the basic result stands unchanged. Our main finding that unemployment, and not inflation, drives interest-rate changes does therefore not seem to be a result of the financial crisis.

4. Discussion

The main objective of the ECB is price stability, and it has succeeded in achieving this fairly well. Our results indicate that it is not caused by adherence to anything remotely close to a Taylor rule during the past 11 years. This fits well with several ECB statements (cf. Asso *et al.*, 2010). Upon reflection, it should not be a cause of concern either.

First, it is well established that what a central bank responds do not reveal its ultimate motives. It may respond to certain variables because they are good indicators for its goal variable(s). Such intermediate targeting is clearly what has driven the ECB's behavior.[8]

[8] This is also emphasized by Gerlach (2007). He considers the ECB's actual, and discrete, decisions on changes in the repo rate in contrast with our focus on a target for the market-determined EONIA. Nevertheless, our focus on interest rate *changes* makes our paper more related to his than to the large level-estimation based literature. Interestingly, and in accordance with our findings, Gerlach's ordered probit analysis does not detect a notable influence of inflation *per se* on interest-rate decisions.

Second, the failure of identifying an adherence to an active Taylor rule should not lead to inference about potential stability problems in the Euro area. As one of us has emphasized elsewhere, empirical estimates are characterizations of equilibrium relationships between macroeconomic variables. If these variables result from a central bank conducting optimal monetary policy in a world that is (somewhat) forward looking, these equilibrium relationships may exhibit small correlations between the nominal interest rate and inflation. Jensen (2009) presents estimations on simulated data from a small-scale New-Keynesian model with optimal discretionary policymaking in a stable and fundamental-based equilibrium. There, data looks as if an inactive Taylor rule has been followed. Furthermore, if optimal policy is conducted under commitment, the central bank's ability to affect expectations causes the correlation between the nominal interest rate and inflation to vanish or even become negative in data at some horizon of inflation expectations. Intuitively, if a central bank can fight current inflation by "talking" inflation expectations down, the current interest rate needs to be increased very little, and its equilibrium correlation with inflation expectations may become negative.

The lack of empirical identification of an active Taylor rule could therefore be seen as an empirical sign of a credible inflation-stabilizing central bank.

Appendix

A. Further details on Table 3

The instrument set for the basic equation is $\mathbf{z}_t^u = \{1, \Delta i_{t-1}, \Delta i_{t-2}, \Delta i_{t-3}, \Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta u_{t-2}, \Delta u_{t-3}, \Delta u_{t-4}, \Delta u_{t-5}\}$. Let Δx_t denote the first difference of an additional variable. Then, the instrument sets for the estimations with additional variables are

Additional variable	$\mathbf{z}_t$
$M3$ growth	$\mathbf{z}_t^u \cup \{\Delta x_{t-2}, \Delta x_{t-3}, \Delta x_{t-4}\}$
Oil price	$\mathbf{z}_t^u \cup \{\Delta x_t, \Delta x_{t-1}\}$
Dollar/Euro exchange rate	$\mathbf{z}_t^u \cup \{\Delta x_{t-1}\}$
Real effective exchange rate	$\mathbf{z}_t^u \cup \{\Delta x_{t-1}\}$
Long Euro-bond rate	$\mathbf{z}_t^u \cup \{\Delta x_{t-1,} \Delta x_{t-2}\}$

In the cases of the Economic sentiment indicator and Lehman Brothers dummies, $\mathbf{z}_t = \mathbf{z}_t^u \cup x_t$. In each equation without HICP inflation, $\mathbf{z}_t^u$ is replaced by $\mathbf{z}_t^u \setminus \{\Delta\pi_{t-1}, \Delta\pi_{t-2}\}$.

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What Drives the European Central Bank's Interest-Rate Changes?

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Supplementary appendices: Not intended for publication

B. Detailed data description

All data are available at `www.ecb.int/stats/`. The details of each series, and our treatments, are:

EONIA:

Official description: Euro area (changing composition) - Money Market - Eonia rate - Historical close, average of observations through period - Euro, provided by ECB. Percent per annum

Official acronym: FM.M.U2.EUR.4F.MM.EONIA.HSTA

Our transformation of the raw data: None

HICP inflation:

Official description: Euro area (changing composition) - HICP - Overall index, Annual rate of change, Eurostat. Neither seasonally nor working day adjusted

Official acronym: ICP.M.U2.N.000000.4.ANR

Our transformation of the raw data: None

Output:

Official description: Euro area 16 (fixed composition) - Gross domestic product at market price - Chain linked volumes, reference year 2000 - ECU/euro - Seasonally and partly working day adjusted, mixed method of adjustment

Official acronym: ESA.Q.I5.S.0000.B1QG00.1000.TTTT.L.U.A

Our transformation of the raw data: Logarithm of quarterly data converted to monthly data by cubic spline transformation

Unemployment:

Official description: Euro area 16 (fixed composition) - Standardised unemployment, Rate, Total (all ages), Total (male & female), Eurostat, Seasonally adjusted, not working day adjusted, percentage of civilian workforce

Official acronym: STS.M.I5.S.UNEH.RTT000.4.000

Our transformation of the raw data: None

M3 growth:

Official description: Euro area (changing composition), Index of Notional Stocks, MFIs, central government and post office giro institutions reporting sector - Monetary aggregate M3, All currencies combined - Euro area (changing composition) counterpart, Non-MFIs excluding central government sector, Annual growth rate. Working day and seasonally adjusted

Official acronym: BSI.M.U2.Y.V.M30.X.I.U2.2300.Z01.A

Our transformation of the raw data: None

Oil prices:

Official description: Brent crude oil 1-month Forward - free on board in US Dollar per barrels; Financial market data type: Historical close, average of observations through period

Official acronym: FM.M.U2.EUR.4F.CY.OILBRNI.HSTA

Our transformation of the raw data: 100 times the logarithm.

Economic sentiment indicator:

Official description: Euro area 16 (fixed composition), EU Commission, DG-ECFIN, Economic sentiment indicator, Total, Seasonally adjusted, not working day adjusted

Official acronym: SUR.M.I5.S.ECFIN.ESI000.TT

Our transformation of the raw data: 100 times the logarithm.

Dollar/Euro exchange rate:

Official description: ECB reference exchange rate, US dollar/Euro

Official acronym: EXR.M.USD.EUR.SP00.A

Our transformation of the raw data: 100 times the logarithm.

Real effective exchange rate:

Official description: ECB Real effective exchange rate CPI deflated, Euro area-16 countries vis-à-vis the EER-41 group of trading partners against Euro

Official acronym: EXR.M.Z60.EUR.ERC0.A

Our transformation of the raw data: 100 times the logarithm.

Long Euro-bond rate:

Official description: Euro area (changing composition) - Benchmark bond - Euro area 10-year Government Benchmark bond yield - Yield - Euro, provided by ECB. Percent per annum

Official acronym: FM.M.U2.EUR.4F.BB.U2_10Y.YLD

Our transformation of the raw data: None

C. Stationarity tests

Table 4: Augmented Dickey-Fuller test statistics

	Data in levels		Data in first differences	
	AR	ARD	AR	ARD
EONIA interest rate	–0.961	–1.834	–4.153‡	–4.147‡
HICP-inflation	–0.865	–2.516	–9.249‡	–9.215‡
Output gap[a]/growth	–2.532†	–2.571	–3.031‡	–3.018†
Unemployment	0.057	–2.464	–2.628‡	–2.618
$M3$ growth	–0.965	–1.686	–3.529‡	–3.553‡
Oil price[b,c]	0.795	–1.474	–3.116‡	–3.112†
Economic sentiments index	–0.154	–2.739	–4.157‡	–4.139‡
Dollar/Euro exchange rate[b,c]	–0.432	–1.547	–1.958†	–2.254
Real effective exchange rate[b]	0.141	–2.383	–8.878‡	–8.847‡
Long Euro-bond rate	–0.294	–1.988	–9.106‡	–9.071‡

Notes: "AR" denotes autoregressive model, "ARD" denotes autoregressive model with drift. In each instance, the Bayesian Schwarz Information Criterion is used in determining number of lags (18 is maximum number of lags).

[a]For the level case, we use an output-gap measure constructed as HP-filtered GDP.

[b]The Bayesian Schwarz Information Criterion chose the longest lag length for the variable in levels.

[c]The Bayesian Schwartz Information Criterion chose the longest lag length for the variable in first differences. For lags up to at least 10, nonstationarity was strongly rejected.

†Rejects nonstationarity at the 5% significance level. ‡Rejects nonstationarity at the 1% significance level.