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ORIGINAL ARTICLE

The influence of government ideology on monetary policy: New cross-country evidence based on dynamic heterogeneous panels

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Abstract

Using data of 23 OECD countries over the 1980–2005 period, we examine whether government ideology affects monetary policy, conditional on central bank independence. Unlike previous studies in this line of literature, we estimate central bank behavior using forward-looking and real-time data in Taylor rule models and use estimators that allow for heterogeneity across countries. Our models with heterogeneous slope coefficients for the full sample do not suggest partisan effects. We also do not find evidence that central bank behavior is conditioned by the interaction of the ideology of the incumbent government and the electoral calendar.

JEL CLASSIFICATION
E52, E58, D72, C23

KEYWORDS
Central Bank Independence, ideology, monetary policy, partisan theory

1 | INTRODUCTION

Do central banks follow policies that are in line with the ideological preferences of the government? According to Goodman (1992: 4–5): “Political parties that win elections are often said to favor those economic policies which reflect the economic interests of their core constituencies. According to this view, conservative parties will pursue more restrictive policies to achieve lower rates of inflation, while left-wing parties will pursue more expansionary policies to achieve lower rates of unemployment.[1] "… In countries with dependent central banks, governments play a significant role in setting monetary policy. As a consequence, monetary policy can be influenced by … the identity of the party in control of government…. In countries with independent central banks, on the other hand, governments have, by definition, little influence over the formulation of monetary policy. … All other things being equal, independent central banks are both more able and more likely than are dependent central banks to resist any domestic political pressure for greater monetary expansion."[2]

Even though central banks in many countries have been designed to minimize political influences on monetary policy decision-making, these protections are generally not perfect so that government ideology may impact central banks' policies (Chappell, 2004). Furthermore, politicians can change legislation on central bank independence (CBI). In the United States, for instance, senator Rand Paul recently reintroduced his proposal to let the Government Accountability Office review the Federal Reserve's monetary policy. The Fed strongly opposes the bill, arguing that it would subject monetary policy-makers to excessive political pressure. However, the bill may become a reality as Republicans, who are overwhelming critical of the Federal Reserve, now control both the House and the Senate.

Several studies have examined the influence of the U.S. administration on the policies of the Federal Reserve. In his seminal contribution, Havrilesky (1995) argues that presidential influence over monetary policy derives in part from the Fed's need to deflect Congressional threats to its independence. In return for protection from the president, the Fed will accommodate presidential wishes. The evidence of Havrilesky (1988, 1995) suggests that the Fed was responsive to executive branch signals of monetary policy preferences. An alternative channel through which politicians may influence monetary policy is through the power of appointment (Chang, 2001; Chappell, Havrilesky, & McGregor, 1993; Havrilesky & Gildea, 1991; and Schnakenberg, Turner, & Uribe-McGuire, 2017). By appointing political allies in the Federal Reserve Board, the president may try to redirect monetary policy in line with his preferences.

In contrast to the prediction of partisan theory, several studies report evidence that short-term interest rates tend to be higher during left-wing governments (Boix, 2000; Clark, 2003; Sakamoto, 2008). Boix (2000) and Clark (2003) suggest that left-wing governments follow a tighter monetary policy in order to convince investors or public opinion of their commitment to fight inflation. By following atypical policies, the government aims to earn credibility in line with the "When Does it Take a Nixon

[1]Goodman refers here to partisan theory (Hibbs, 1977). According to this theory, right-wing governments care more (less) about inflation (unemployment) than left-wing governments, in line with the interests of their respective constituencies. Within a rational expectations framework with optimizing subjects and in the context of election uncertainty, differences between left- and right-wing governments will only be temporary and show up shortly after the elections (Alesina, 1987; Alesina & Sachs, 1988). Of course, right-wing and left-wing governments may differ along several other dimensions as well, like taxes and regulation. Potrafke (2017) provides a comprehensive review of the empirical evidence on partisan politics in OECD panel studies. See also Cahan, Lorenz, and Potrafke (2017) for an extensive discussion of panel and single-country studies on the impact of ideology on monetary policy. Although several central banks in our sample have become inflation targeters, this does not imply that partisan factors have become irrelevant; the reason is that inflation targeting does not imply that central banks only care about inflation (Samarina & de Haan, 2014).

[2]As argued by de Haan, Bodea, Hicks, and Eijffinger (2018), the economic case for CBI rests on countering inflationary biases that may occur for various reasons in the absence of an independent central bank. One reason for such a bias is political pressure to boost output in the short run for electoral reasons. Another reason is the incentive for politicians to use the central bank's power to issue money as a means to finance government spending. The inflationary bias can also result from the time-inconsistency problem of monetary policy-making. In a nutshell, this is the problem that policy-makers are not credible, that is, they have an incentive to renege in the future on their promise made today to keep inflation low. As people are aware of this temptation, they have higher inflation expectations, which would lead to higher inflation without any gains in employment or output. By delegating monetary policy to an independent and conservative (i.e., inflation averse) central bank, promises to keep inflation low are more credible.

to Go to China" idea of Cukierman and Tommasi (1998). Sakamoto (2008) provides another explanation, arguing that lack of control over the central bank incentivizes left-wing governments to use expansionary fiscal policy, which, in turn, forces independent central banks to increase interest rates.

Belke and Potrafke (2012) estimated a dynamic fixed effects panel model using quarterly data over 1980–2005 of 23 developed countries. Their analysis is based on estimates of (backward-looking) Taylor rules (Taylor, 1993), incorporating central bank independence and government ideology. Belke and Potrafke (2012) provide support for traditional partisan effects conditional on CBI: short-term interest rates are lower during a left-wing government when the central bank is not independent. However, under independent central banks and left-wing governments, interest rates are higher.

Clark and Arel-Bundock (2013) argue that conservative central bankers behave as conditional inflation hawks because doing so can result in the (re)election of political principals that more closely share their low-inflation preferences. Their evidence for the United States, based on estimates of a (partly forward-looking) Taylor rule, suggests that when Democrats control the White House the Fed increases interest rates in response to increased inflationary expectations, especially as elections approach, but it is insensitive to higher inflationary expectations under a Republican administration.

Several previous studies (including Belke & Potrafke, 2012 and Cahan et al., 2017) assume that monetary authorities only consider backward-looking information about output and inflation when taking policy decisions. Indeed, the standard procedure in macroeconomic and political science literature on modeling economic policies has been to use the latest available data. However, as pointed out by Croushore (2011), this approach is based on the heroic assumptions that data are immediately available (when in fact they are generally available only with a lag) and that data revisions either do not exist or are inconsequentially small (when in fact they are often large and may significantly affect empirical results).[3] Indeed, the main reason why estimates of the Taylor rule based on ex post, backward-looking data deviate from those based on real-time, forward-looking data is revisions in the estimates of the output gap.[4] In addition, monetary policy instruments (such as the policy interest rate) only affect central bank target variables (such as output and inflation) with substantial delay, making the use of backward-looking information suboptimal. Several studies have shown that central banks are forward looking (see, for instance, Clarida, Galí, & Gertler, 1998; Orphanides, 2001; Svensson, 2003; Gorter et al., 2008; and Blinder, Ehrmann, de Haan, Fratzscher, & Jansen, 2008).

This paper examines the influence of governments' ideology on the actions of central banks for a large sample of OECD countries, conditional on the level of central bank independence. Our first contribution is that, unlike previous studies which mostly used backward-looking and ex post data in modeling central bank behavior, our preferred models are based on real time, forward looking data for the reasons outlined above. These forward-looking data have not been subject to posterior adjustments and/or revisions. We estimate panel models using data of 23 OECD countries over the 1980–2005 period for several reasons. First, real-time data are not available for a sufficiently long period for most countries and, second, there is not sufficient variation in partisan factors and CBI over time to identify conditional partisan influences in single-country models. Indeed, for the latter reason several previous

[3]This is nicely illustrated by Orphanides (2001) who finds that estimated policy reaction functions for the Federal Reserve based on ex post data yield misleading descriptions of the Fed's policy-making. Models based on real-time Federal Reserve staff forecasts describe policy better than comparable Taylor-type specifications based on ex post data. This is further exemplified by the work of Gorter, Jacobs, and de Haan (2008), who report similar results for the ECB's policies.

[4]An interesting analysis of data revisions is the study by De Castro, Perez, and Rodríguez-Vives (2013) who examine revisions in government budget deficit data for European Union member states. They find that initially released government deficit data are biased and nonefficient predictors of subsequent releases, with later vintages of data tending to show larger deficits on average. Furthermore, their results suggest that governments tend to conceal deficits in electoral and pre-electoral years.

studies also used panel data. By focusing on the sample period 1980–2005, we make our analysis comparable with previous research, notably the study of Belke and Potrafke (2012).[5]

However, the use of panel models may cause some problems, the most important one being that it may not be correct to assume that the same data-generating process exists in all countries, especially when countries with dissimilar economic and institutional characteristics as Greece and Switzerland are included in the sample (Pesaran & Smith, 1995).[6] This is where the second major contribution of our paper comes in. We employ techniques that allow taking this heterogeneity across countries into account. Our results cast serious doubts on the presence of partisan effects in monetary policy. When the full sample is used and heterogeneous slope coefficients are allowed for, we do not find an association between government ideology and monetary policy.

Our third contribution is that we follow a similar approach as Clark and Arel-Bundock (2013) and examine whether conservative central bankers adjust interest rates strategically before elections. Our results do not provide strong evidence for this view neither.

The remainder of the paper is structured as follows. Section 2 presents a brief description of the data used. Section 3 describes the empirical approach. Section 4 shows the findings for models similar to those of Belke and Potrafke (2012) which include interactions between CBI and government ideology. Section 5 tests the hypothesis of Clark and Arel-Bundock (2013) that central banks react differently to inflation expectations under left- and right-wing governments, again conditioning for CBI. Section 6 offers the conclusions.

2 | DATA

The dependent variable is the short-term interest rate, which in most cases consists of the annualized 3-month contemporaneous interest rate, and is obtained from the OECD's online database.[7] This also holds for other economic data. Forward-looking data, that is, forecasts for annualized real GDP growth and inflation, have been collected from Consensus Economics. As one of the world's leading economic survey organizations in the world, this firm surveys more than 700 economists and institutions each month to obtain its predictions. This source provides a more consistent and methodologically homogeneous set of information than using each central bank's own forecasts, which in many cases are not available for the time period and frequency considered in our study. Central banks frequently take these forecasts into account when taking monetary policy decisions (Gorter et al., 2008). We therefore use these data as proxy for real-time central bank forecasts. Consensus Economics provides monthly figures. Following the methodology presented in Gorter et al. (2008), these figures are transformed as follows. For any month m of a given year t, the 1-year forecast for a given variable is computed as $(13 - m)/12$ times the forecast for year t plus $(m - 1)/12$ times the forecast for year $t + 1$.

[5]In addition, focusing on this sample period avoids incorporating the noise of the 2007/8 financial crisis and the period of the Effective Lower Bound (ELB) in the dataset. Furthermore, it also minimizes the degree to which the panel is unbalanced due to the inclusion of countries in the European Economic and Monetary Union (EMU).

[6]Cahan et al. (2017) go even so far as to argue that panel models for OECD countries do not identify causal effects of government ideology on monetary policies, and there is no suitable research design to do so. Therefore, the authors survey the historical evidence on each country. Our position is that OECD panel models may be useful for the issue at hand, provided properly specified central bank reaction functions are estimated and heterogeneity is considered.

[7]Available at http://stats.oecd.org. We use money market rates instead of policy rates as the latter not only infrequently change but also because monetary policy is more than changes in policy rates. Notably open market operations are an integral part of monetary policy and their impact is reflected in money market rates. Having said that, also other factors than monetary policy may affect money market rates, but it is widely believed that the main driver of changes in the money market rate is monetary policy.

Subsequently, these values have been averaged into quarterly figures to match the periodicity of other data used. Forward-looking Taylor rules normally contain forecasts of the future output gap (cf. Svensson, 1997, 2003). However, Consensus Economics does not publish such forecasts. We therefore follow the approach suggested by Gorter et al. (2008) to obtain a proxy for this forward-looking variable. Trend GDP growth rate is first estimated for each country, and then this value is subtracted from the expected GDP growth rate to obtain the output gap.[8] In our analysis we also use ex post data to compare our findings with those of previous studies. The ex post output gap is obtained by applying the Hodrick–Prescott filter to real output.[9] Inflation is proxied by the change in the consumer price index (CPI).

The index of government ideology is from Potrafke (2009).[10] This index categorizes countries' governments on a right-to-left scale that ranges from 1, indicating the presence of a strong right-wing ideology, to a value of 5, showing a strong left-wing administration. If the country's government consists of a centrist party, or a balanced coalition of both right- and left-wing parties, the index takes a value of 3. For periods with alternating governments with different ideologies, the index has a value according to the ideology of the government that spent more days in office. Data on CBI come from Arnone, Laurens, Sommer, and Segalotto (2007) and the updates thereof provided by Klomp and de Haan (2009).[11] Following Belke and Potrafke (2012), we use a measure of central bank dependence (CBD), which is the inverse of the CBI variable, which ranges between 0 (low dependence) and 1 (high dependence). Furthermore, whenever the analysis incorporates simultaneously the variables CBD, ideology, and the interaction term between these two, all variables are normalized with a mean of zero and variance of one to allow direct interpretation of the resulting coefficients.

The quarterly backward-looking data used to perform a poolability (homogeneity) test is available for 23 developed countries covering the years of 1980Q4 to 2005Q4. The countries in our sample are: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, and the United States. However, for a group of countries specific events limit the use of all the historical information available: for Germany, the data considered only covers the period after Reunification (1991Q1), and for countries that joined the European Economic and Monetary Union (EMU) the information is restricted to the period in which these countries still had their own monetary policy.[12] In the analysis based on forward-looking data, the sample is reduced to 21 countries (there are no Consensus forecasts available for Iceland and Luxembourg). The sample periods are also smaller. Forecasts for inflation and real GDP growth in Greece are only available from 1993Q2, and forecasts for both Australia and New Zealand are not published in the original source after 1994Q3. In most cases, the quarterly forecasts range from 1989Q4 to 2005Q4. Descriptive statistics of variables included in the backward- and forward-looking models are presented in Appendix A.[13]

[8]Trend GDP growth is the geometric average obtained from the historical real GDP growth series (based on annual data retrieved from the OECD's database for 1980–2005).

[9]Lambda used = 1600.

[10]The data were generously provided by Niklas Potrafke.

[11]Table A3 in Appendix A shows that in our sample CBI varies substantially, both across countries and over time. de Haan et al. (2018) show that even though CBI increased notably in most countries during the 1990s, CBI differs substantially across countries.

[12]Austria, Belgium, Finland, France, Germany, Ireland, Italy, The Netherlands, Portugal, and Spain joined the EMU in 1999. Greece became a member in 2001. For Greece we only consider data until the last quarter of 2000.

[13]We use the Im-Pesaran-Shin (IPS) test and Fisher-type tests to investigate the presence of unit roots and Westerlund's error correction-based tests to examine panel cointegration. The results suggest the presence of a long-run relationship between the variables considered. These results are available in the Supporting information Appendix S1.

3 | EMPIRICAL APPROACH

Central bank policies are represented by a Taylor rule, which models the behavior of short-term nominal interest rates by taking the output gap and the rate of inflation into account (Taylor, 1993).[14] The model also includes a one-period lag of the short-term interest rate as explanatory variable to reflect the interest rate smoothing behavior of central banks (English, Nelson, & Sack, 2002). In addition, the model incorporates variables that represent government ideology, central bank dependence (CBD), and the interaction between these last two terms (cf. Belke & Potrafke, 2012).

Following previous studies, we first assume a backward-looking central bank. The following model is estimated:

$$R_{it} = \beta_1 R_{i,t-1} + \beta_2 ID_{it} + \beta_3 CBD_{it} + \beta_4 \Pi_{\mathrm{it}} + \beta_5 Y_{it} + \beta_6 (ID \times CBD)_{it} + \eta_t + \mu_i + \varepsilon_{it}, \tag{1}$$

where, R_{it} is the short-term nominal interest rate in country i (i = 1, 2, …, N) during period t (t = 1, 2, …, T), while ID_{it} measures the political ideology of the government, *CBD* is the index for central bank dependence, Π_{it} is inflation, and Y_{it} is an indicator of the output gap. The model includes country fixed effects, μ_i, to capture individual unobserved time-invariant factors that influence interest rates. Time fixed effects, η_t, are also part of the model. The disturbances ε_{it} are assumed to be independently distributed across groups.

Next, we estimate a Taylor rule model under the assumption of a forward-looking central bank. In this case the model is:

$$R_{it} = \beta_1 R_{i,t-1} + \beta_2 ID_{it} + \beta_3 CBD_{it} + \beta_4 E(\Pi)_{it} + \beta_5 E(Y)_{it} + \beta_6 (ID \times CBD)_{it} + \eta_t + \mu_i + \varepsilon_{it}, \tag{2}$$

where, $E(\Pi)_{it}$ and $E(Y)_{it}$ are the 12-months ahead forecasts of inflation and the output gap available at time t, respectively. The other variables have the same meaning as in Equation 1.

The presence of a partisan effect on monetary policy implies a negative marginal effect of ideology, but this effect may be conditional on the level of *CBD*. The expected sign for the slope coefficients β_4 and β_5 is positive, while the smoothing parameter (β_1) is expected to be positive and smaller than one.

Due to the presence of a lagged dependent variable (R_{it-1}) as regressor, traditional panel fixed-effect estimators are biased (Nickell, 1981). When the number of cross-sectional units is small, alternative Generalized Method of Moments (GMM) estimators such as the one proposed by Arellano and Bond (1991) and Blundell and Bond (1998) are also biased. Kiviet (1995) and Judson and Owen (1999) provide evidence in favor of using the least squares dummy variable estimator instead, mainly due to its smaller relative variance. Therefore, we employ Bruno's (2005) least squares dummy variable corrected (LSDVC) estimation routine for dynamic panels (see also Belke & Potrafke, 2012).[15]

[14]In the Taylor rule model, nominal policy rates are assumed to be driven by the long-run equilibrium real (or natural) interest rate, the gap between inflation and the central bank's inflation objective, and the output gap, that is,, the deviation of actual and potential output, expressed as a percentage of the latter. Taylor suggested particular values for the natural rate and the parameters for the inflation and output gap. Several studies have found that in recent years, actual policy rates fell below those implied by this particular formulation of the Taylor rule. Hofmann and Bogdanova (2012) have argued that deviations of policy rates from those implied by the normative Taylor rule can be best interpreted as a change in the global equilibrium real (or natural) interest rate. Several studies suggest that the natural rate has declined (see, for instance, Laubach & Williams, 2015; Belke & Klose, 2017). In our analysis, the Taylor rule model is used as a central bank reaction function, that is, to describe central bank behavior, and we therefore do not impose values for the parameters on the inflation and the output gap but estimate them.

[15]This estimation procedure extends the bias approximation formulas in Bun and Kiviet (2003) to accommodate unbalanced panels with a strictly exogenous selection rule. Flannery and Hankins (2013) show that compared to other estimators, LSDVC appears to be the best choice in case of endogenous regressors.

Past empirical research on the influence of ideology on monetary policy mainly used fixed effect panels or similar pooled estimation methods. However, using pooled estimators when the homogeneity assumption does not hold will result in biased estimates of the slope coefficients. This also holds for the LSDVC estimator. Ignoring cross-country heterogeneity induces serial correlation in the disturbance term. This problem is present even if T goes to infinity (Pesaran & Smith, 1995). When T is sufficiently large so that separate regressions for each country can be run, the homogeneity assumption can be verified using the test proposed by Baltagi (2008) and adapted to a dynamic framework in Baltagi and Griffin (1997). This test is based on the following statistic:

$$F = \frac{(SSE_r - SSE_u/q)}{SSE_u/dfu} \sim F(q, dfu)$$

where SSE_r is the sum of squared residuals of a two-stage least squares (2SLS) regression that assumes common intercepts and common slopes, SSE_u is the sum of squared residuals of a 2SLS regression that allows for both varying intercepts and slopes, and q and dfu are the number of restrictions and the degrees of freedom of the unrestricted model, respectively.

To assess the suitability of pooling, we first perform this test on data for the same period and relationships considered in Belke and Potrafke (2012) as these authors provide the most comprehensive panel estimations using backward-looking data. Their analysis covers 23 OECD countries from 1980 to 2005. The null hypothesis of homogeneity is rejected at the 1% confidence level as shown in Appendix B.

An intuitive approach to deal with the problem of cross-country heterogeneity (i.e., differences in the data-generating process across countries) is to find a subsample of countries that is sufficiently similar. As a first step, we therefore split the sample in countries with above and below-average inflation.[16] Baltagi's poolability test indicates that when dealing with backward-looking information, the group of 16 countries with below-average inflation have homogeneous slope coefficients (Appendix B). These countries are: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Japan, Luxembourg, the Netherlands, Norway, Sweden, Switzerland, United Kingdom, and the United States.

Pesaran and Smith (1995) and Pesaran, Shin, and Smith (1999) propose alternatives to the pooled approach that can deal with cross-country heterogeneity, namely the mean group (MG) estimator and the pooled mean group estimator (PMG).[17] We apply these approaches for our preferred forward-looking specification of the Taylor rule. The MG estimator can be explained as follows. Take the following random coefficient model (RCM):

$$R_{it} = \lambda_i R_{i,t-1} + \beta_i' \boldsymbol{x}_{it} + \eta_i + \varepsilon_{it} \tag{3}$$

where R_{it} is a vector with observations of the short-term interest rate and $\boldsymbol{x}_{it}$ is a vector of exogenous regressors that includes (the forecasts of) inflation and the output gap, and the other variables specified in Equation 2. Notably, λ_i and β_i are varying across groups according to the following specification:

$$\lambda_i = \lambda + \xi_{1i}, \beta_i = \beta + \xi_{2i}, \tag{4}$$

where ξ_{1i} and ξ_{2i} are assumed to have zero means and constant covariances. As such, this is a standard formulation of the RCM, and it introduces heterogeneity in the parameters through the 'short-term'

[16]Using the interest rate variable to split the sample also generates a cluster with homogeneous coefficients, but the number of countries in this group is smaller.

[17]Other estimators for heterogeneous panels, such as the 'shrinkage' estimator suggested by Maddala, Srivastava, and Li (1994), lack the property of distinguishing between short-run and long-run effects. We leave it for future research to consider these alternative estimators.

coefficients λ_i and β_i. If the sample is large enough, then the mean of the long-run effects $\theta_i = \beta_i/(1-\lambda_i)$, denoted $\bar{\theta}$, becomes the appropriate measure of the average long-run effect.

The PMG estimator allows estimating a common long-run relationship without assuming similar dynamics in each country (Pesaran et al., 1999). We follow the approach suggested by these authors, which starts by assuming an autoregressive distributed lag dynamic (ARDL) panel specification, ARDL(p, q_1, q_2, …, q_z), where p refers to the number of lags of R_{it}, q is the number of lags of the independent variables, and z is the number of regressors. In general terms, this specification has the following form:

$$R_{it} = \sum_{j=1}^{p} \lambda_{ij} R_{i,t-j} + \sum_{j=0}^{q} \delta_{it}' \boldsymbol{x}_{i,t-j} + \mu_i + \varepsilon_{it} \tag{5}$$

where $\boldsymbol{x}_{it}$ is a vector of exogenous regressors for group i as defined in Equation 2, the λ_{ij} are scalars and the rest of the variables have the usual interpretation. To keep the notation simple, no time fixed effects are included in Equation 5. If the variables in Equation 2 are cointegrated, then the error term is an I(0) process (stationary) for all countries. This implies that the equation can be written in an error correction form:

$$\Delta R_{it} = \phi_i \left(R_{i,t-j} + \theta_i' \boldsymbol{x}_{it} \right) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta R_{i,t-j} + \sum_{j=0}^{q-1} \delta_{it}^{*\prime} \Delta \boldsymbol{x}_{i,t-j} + \mu_i + \varepsilon_{it} \tag{6}$$

where $\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$, $\theta_i = -\frac{\beta_i}{\phi_i}$, $\beta_i = \sum_{j=0}^{q} \delta_{ij}$, and Δ represents a first difference operator. Also, $\lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im}$ with $j = 1, 2, \ldots, p-1$ and $\delta_{ij}^{*} = -\sum_{m=j+1}^{p} \delta_{im}$ *with* $j = 1, 2, \ldots, q-1$. The parameter ϕ_i is the error-correcting speed and if $\phi_i < 0$ then a long-run relationship between R_{it} and $\boldsymbol{x}_{it}$ exists and is defined for each i by:

$$R_{it} = -\left(\frac{\beta_i'}{\phi_i}\right) \boldsymbol{x}_{it} + \upsilon_{it}$$

where υ_{it} is a stationary process, and the long-run coefficients defined by the θ_i are the same across countries. Pesaran et al. (1999) use a maximum likelihood approach to estimate this model.

Finally, a Hausman test is performed to decide whether the MG or the PMG estimator is to be preferred.[18]

4 | RESULTS: PARTISAN EFFECTS CONDITIONAL ON CENTRAL BANK DEPENDENCE

4.1 | Homogeneous sample

Table 1 presents the estimates of Equations 1 and 2 using Bruno's bias-corrected least squares dummy variables estimator. The first three columns show estimations of Equation 1 using different samples and the fourth column presents the estimation results of Equation 2. In particular, the first column shows the results for a backward-looking monetary policy model similar to models estimated by Belke and Potrafke (2012) for the complete sample of 23 countries. Although the poolability test

[18]The PMG estimator constrains the long-run elasticities to be equal across all panels. This "pooling" across countries yields efficient and consistent estimates when the restrictions are true. If the true model is homogeneous, the PMG estimates are consistent and more efficient. MG estimates are consistent in either case but are only efficient if the true model is heterogeneous (Blackburne & Frank, 2007). If the true model were homogeneous, we could have faced the problem that two sets of consistent estimators would be available. However, in our case, heterogeneity prevails and the MG approach provides the only set of consistent and efficient estimators. More methodological details are available in Appendix E.

(Appendix B) suggests that it is likely that the estimated coefficients of this model are not homogeneous across countries, and are hence inconsistent, we start with this model to serve as a benchmark.[19] The second column shows the results for the backward-looking model using the sample of 16 countries having similar inflation rates and for which the poolability test indicates the presence of homogeneous slope coefficients.[20] The main difference between the first and the second column is therefore the number of countries considered. In column (3) also Luxembourg is dropped, as forward-looking information is not available for this country, and the time period is adjusted to match the characteristics of the forward-looking sample available. The results in column (3) can then be compared with those in column (4), which shows the outcomes for the Taylor rule model using real-time (1-year) forecasts of inflation and the output gap.[21] For the forward-looking model, the poolability test indicates homogeneous slope coefficients at the 1% confidence level (Appendix B).

Similar to the results reported in most other papers discussed in section 1, all models shown in Table 1 deviate from the ‘Taylor principle’ (Taylor, 1993). In a nutshell, this means that the long-term coefficient on inflation should be larger than 1 for monetary policy to have a stabilizing impact on inflation. A potential explanation why this principle does not hold is that the period considered in this analysis largely coincides with the ‘great moderation’, which was a period of low and stable inflation. In that sense, it may just reflect central banks’ success in keeping inflation stable and low over the period considered. In the absence of major movements in inflation, the reaction of interest rates to inflation might simply have become more difficult to pin down with any great precision (Hofmann & Bogdanova, 2012). Alternatively, it might suggest misspecification of the models used (see section 4.2 for further details).

The effect of ideology on the short-term interest rate is, ceteris paribus*,* only significantly different from zero when a forward-looking central bank is considered (first line, last column, in Table 1). However, to measure the complete effect of ideology on the interest rate, the marginal effect has to be computed and evaluated at different levels of CBD. The marginal effects for the homogeneous backward-looking models shown in Table 2 (columns 2 and 3) suggest that interest rates are higher under left-wing administrations when the central is bank is highly independent. These results confirm Sakamoto’s (2008) proposition and Belke and Potrafke’s (2012) finding that—within the context of a backward-looking Taylor rule and homogeneous panel model—interest rates are lower under a left-wing government if the central bank is strongly dependent.

However, if monetary policy is modeled in a forward-looking manner, the results deviate from the findings of previous studies on partisan effects in monetary policy (column 4). At high levels of central bank dependence, no statistically significant effect of ideology on interest rates is found. But at average levels of CBD the effect is positive and it gets larger in size and statistical significance the less dependent the central bank is. At a minimum degree of dependence, an increase in the ideology index of one standard deviation (i.e., the government becomes more leftist) will raise the interest rate by almost 0.2 percentage points. To put this result in perspective, if the control of the government changes from a right-wing to a left-wing party after elections, interest rates will be expected to rise more than 0.4 percentage points on average. Due to the homogeneity assumption, the absolute size of this change is expected to

[19]Appendix C confirms that the results of the model shown in column (1) of Table 1 are sensitive to the exclusion of countries. Belke and Potrafke (2012) report similar results.

[20]Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Japan, Luxembourg, the Netherlands, Norway, Sweden, Switzerland, United Kingdom, and the United States.

[21]Even when considering the same periods and countries, differences in the availability of backward and forward-looking data explain the difference in the number of observations between columns 3 and 4. Further adjustments to the sample were made in order to obtain exactly the same number of observations as in Belke and Potrafke (2012). This does not change our conclusions as shown in Appendix D.

TABLE 1 The impact of government ideology on monetary policy

	Backward-looking I (heterogeneous)	Backward-looking II (homogeneous)	Backward-looking III (homogeneous)	Forward-looking (homogeneous)
Ideology	−0.018	−0.007	0.015	0.064**
Central Bank Dependence (CBD)	0.164***	0.120***	0.162***	0.151***
CBD*Ideology	−0.063**	−0.096***	−0.089***	−0.079***
(Expected) Output gap	0.122***	0.141***	0.139***	0.058*
(Expected) Inflation	0.111***	0.072***	0.022	0.068
Lagged Dep. Var.	0.786***	0.855***	0.882***	0.883***
Fixed country effects	Y	Y	Y	Y
Fixed time effects	Y	Y	Y	Y
Period	1980/2005	1980/2005	1989/2005	1989/2005
Observations	1,884	1,383	787	742
Number of countries	23	16	15	15

Notes. Dependent variable: short-term interest rate (annualized quarterly percentage). CBD is central bank dependence (ranging from 0 to 1), Ideology ranges from 1(strong right) to 5 (strong left), Output gap is the annualized percentage long-term GDP growth gap, and Inflation is the annualized quarterly consumer inflation rate. Regressions estimated using Bruno's bias-corrected least squares dummy variables estimator. Ideology and central bank dependence are interacted (normalized). Statistical significance: *significant at 10%; **at 5%; ***at 1%.

be the same in each country. Countries with lower average interest rates will suffer larger proportional changes than countries with historically higher interest rates. This result is rather implausible.

4.2 | Forward-looking heterogeneous sample

An alternative to consider only subsamples of countries with homogeneous coefficients can be to use estimators that allow for nonhomogeneous slope coefficients. The mean group (MG) estimator assumes that the slope coefficients are different across all cross sections. The MG estimation is conducted using an error correction specification (Equation 6). The pooled mean group (PMG) estimator imposes homogeneity on the slope coefficients of the long-run relationship only, while allowing for short-run heterogeneity.

TABLE 2 Marginal effect of ideology on interest rates for different levels of central bank dependence

	(1) Backward-looking I (heterogeneous)	(2) Backward-looking II (homogeneous)	(3) Backward-looking III (homogeneous)	(4) Forward looking (homogeneous)
Minimum CBD	0.117 (0.078)	0.182*** (0.065)	0.165*** (0.064)	0.192*** (0.045)
Average CBD	−0.018 (0.035)	−0.008 (0.029)	0.015 (0.027)	0.062* (0.032)
Maximum CBD	−0.109* (0.055)	−0.160** (0.066)	−0.147** (0.063)	−0.078 (0.054)

Notes. See notes to Table 1. The table shows the marginal effect of ideology of the regressions shown in Table 1. Standard errors in parentheses. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

This approach has not been used before to examine the relationship between ideology and monetary policy, but has been employed in other lines of research (Klomp & de Haan, 2013; Pesaran et al., 1999). Since in reality central banks use forecasts to take decisions, this part of the analysis focuses on forward-looking Taylor rule models. As explained in the previous section, the PMG estimation corresponds to an ARDL specification. We limit the number of lags in this specification to a minimum. This is not only closer to a Taylor rule than other specifications of ARDL models but also allows limiting the number of parameters to be estimated. A common lag structure across countries is also imposed. Table 3 presents the results for the long-run coefficients only.[22]

For each estimation, the long-term coefficients as shown in the first column result from the estimation of the error correction models, while the second column presents the *z*-score to determine how statistically significantly different from zero each coefficient is. The parameter ϕ, reflecting the speed of adjustment, is also included to illustrate the speed at which the model converges to the long-run equilibrium. Imposing homogeneity on the long-run coefficients creates an upward bias in the size of the coefficient of the lagged dependent variable $R_{i,t-1}$ (Pesaran et al., 1999). As a consequence, the MG estimate of the adjustment speed is larger than that of the PMG approach.[23] Another consequence of imposing this long-run homogeneity is a reduction in the standard errors of the long-run coefficients.

Under the assumption of long-term homogeneity in the slope coefficients, the coefficients of expected inflation and output have the expected sign and are significantly different from zero at the 1% level. However, the marginal effect of government ideology conditional on CBD is only significantly different from zero and positive at a minimum level of CDB (Table 4). In other words, if it assumed

TABLE 3 The impact of ideology on forward-looking monetary policy

	PMG		**MG**	
	Coefficient	***z*-Value**	**Coefficient**	***z*-Value**
Long-run relation				
Ideology	0.893***	4.04	−0.396	−0.4
Central Bank Dependence (CBD)	0.165	0.41	0.184	0.7
CBD*Ideology	−0.535*	−1.87	−2.517	−1.20
(Expected) Output gap	3.223***	5.29	1.617*	1.88
(Expected) Inflation	2.628***	11.2	1.872**	2.24
Error correction speed	−0.090***	−6.59	−0.214***	−5.55
Number of *N*	21		21	
Observations	935		935	
Period	1989/2005		1989/2005	

Notes. Heterogeneous panel. Error correction model dependent variable: short-term interest rate (annualized quarterly percentage). CBD is central bank dependence (ranging from 0 to 1), Ideology ranges from 1 (strong right) to 5 (strong left), Output gap is the annualized percentage long-term GDP growth gap, and Inflation is the annualized quarterly consumer inflation rate. Pesaran's Pooled Mean Group and Mean Group estimators are obtained in STATA using *xtpmg* command. If ideology and central bank dependence interact, the data are normalized. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

[22]Stata command *xtpmg* is used, with options "p*mg*" and "*mg*" in each case. Table E1 shows the full model estimates.

[23]Recall that imposing restrictions on parameters leads to a bias unless the restrictions are true. Also, remember that $\theta_i = -\frac{\beta_i}{\phi_i}$, and that $\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$.

TABLE 4 Marginal effects of ideology on interest rate conditional on central bank dependence allowing for heterogeneity

	PMG	MG
Minimum CBD	1.835*** (0.545)	4.039 (4.245)
Average CBD	0.912 (0.221)	−0.305 (1.031)
Maximum CBD	−0.005 (0.532)	−4.626 (3.213)

Notes. See notes to Table 3. Standard errors in parentheses. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

that the slope coefficients are homogeneous in the long term, interest rates are higher under left-wing administrations but only if the forward-looking central bank is highly independent.

However, under the assumption of heterogeneity in the slope coefficients, only the two traditional components of the Taylor rule are different from zero (MG). The impact of expected inflation on the interest rate is larger than one, in line with the Taylor principle. When analyzed independently, ideology and central bank dependence do no enter this model with coefficients statistically different from zero. On average, when all countries have different slope coefficients, no partisan effect is present as shown in the second column of Table 4.

As the results of the PMG and the MG estimators clearly differ, it is important to decide which estimator should be preferred. Therefore, a Hausman specification test on the null hypothesis of homogeneous slope coefficients is performed (Hausman, 1978). This test compares an estimator b that is known to be consistent with an estimator B that is efficient under the assumption being tested. The null hypothesis is that the estimator is an efficient (and consistent) estimator of the true parameters, and its test statistic is Chi-squared distributed with degrees of freedom equal to the number of slope coefficients being compared. Under the null hypothesis of homogeneity, both the mean group and the pooled mean group estimators are expected to provide asymptotically consistent estimates. In turn, the PMG estimates will be inconsistent under the alternative hypothesis of heterogeneity. The test results suggest that the null is rejected, albeit only at the 10% level, which implies that all slope coefficients are heterogeneous across countries and the MG estimator is to be preferred.[24] In light of this evidence, we conclude that our preferred estimates do not suggest the presence of a partisan effect in monetary policy.

5 | STRATEGIC CENTRAL BANKS?

Clark and Arel-Bundock (2013) argue that there are systematic differences in Fed behavior under Republican and Democratic presidents since the U.S. central bank became operationally independent in 1951. Their evidence suggests that the Fed's policy instrument, the Federal Funds Rate, responds to the inflation rate only when Democrats are in the White House and up for re-election: The Federal Reserve is a conditional inflation hawk, that is, it cares about inflation, but only when the President is a Democrat and in the close vicinity of an election date. The underlying theoretical reasoning is that the preferences of the Republic Party are closer to those of the Fed than those of the Democratic Party. As a consequence, the central bank has an incentive to behave strategically: "central bankers in such a world face an intertemporal tradeoff when right-wing governments are in power. They can push for the policy closest to their ideal point knowing that doing so hinders the right-wing party's prospects of reelection, or they can accept a more expansionary policy now in exchange for an increase in the

[24]Test statistic of 9.31 with a p-value equal to 0.097.

TABLE 5 Marginal effects of ideology on interest rate conditional on election date allowing for heterogeneity

	Countries with low CBD	Countries with high CBD
After election (Before_elec=0)	0.135 (0.46)	−0.526 (0.88)
Before election (Before_elec=1)	−1.673 (1.15)	−0.086 (1.51)

Notes. See notes to Table 3. Interacted variables are normalized. Standard errors in parentheses. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

probability that the government in power during non-electoral times will be closer to its ideal point. In contrast, when left-wing parties are in power, a conservative central banker faces no such dilemma. Political preferences align with the Fed's well publicized commitment to price stability, since the political consequence of following through on that commitment is an increase in the probability that a government with an ideal point closer to its own will get elected. Thus, we can expect independent central banks to act like conservative independent central banks—but only when left-wing governments are in power." (Clark & Arel-Bundock, 2013, p. 5).

Based on the findings of the previous sections, and within the framework of our multicountry analysis, we test this hypothesis by computing the mean group estimators of our preferred forward-looking Taylor rule for two subsamples, namely countries above and below the median level of CBD. We introduce a new dummy variable (Before_elec) which equals one for periods before the main elections in each country of our sample.[25] This new variable and an interaction between it and government ideology are added as explanatory variables to our initial model.[26] We have also normalized the variables included in the interaction terms to facilitate the interpretation of the results. The hypothesis that the effect of government ideology on interest rates being conditional on the electoral calendar implies a positive marginal effect on higher values of government ideology before the election (i.e., under left-wing governments), but only when the central bank is not dependent (conservative). The results of this new specification of the model are presented in Table 5 and no marginal effect is statistically significantly different from zero at any level of confidence. Our estimates do not provide support for this hypothesis.

As a robustness check, we also run an alternative specification of the model in which we do not split the sample and incorporate a triple interaction term between the variables CBD, government ideology, and the election variable.[27] Considering three different levels of CBD in this new specification (minimum, average, and maximum), we still find that the marginal effects of ideology on the interest rate conditional on our election variable are not significantly different from zero (Table 6).[28]

6 | CONCLUDING REMARKS

Our results do not provide robust evidence for the presence of a traditional partisan effect on monetary policy. In contrast to the standard assumption used in previous research, our evidence suggests that the influence of government ideology on the interest rate cannot be assumed to be equal across large samples of countries. The preferred estimation method in the literature, fixed-effect panels, can therefore only be used for smaller samples of similar countries without creating any biases, but even

[25]The methodology for constructing this variable is presented in Appendix G.

[26]See Equation 2 for the base model. The terms CBD and ID × CBD are not included in the estimates shown in Table 5, but a regression including these terms provides similar conclusions.

[27]We thank an anonymous referee for this suggestion.

[28]Table F1 and F2 in Appendix F present the complete estimation results.

TABLE 6 Marginal effects of ideology on interest rate conditional on election date, CBD, and allowing for heterogeneity

	Minimum CBD	Average CBD	Maximum CBD
After election (Before_elec=0)	12.02	1.334	−9.29
	(11.75)	(3.06)	(6.32)
Before election (Before_elec=1)	8.521	2.185	−4.116
	(7.82)	(1.68)	(5.05)

Notes. See notes to Table 3. Interacted variables are normalized. Standard errors in parentheses.

in that case the policy of a forward-looking central bank is not affected as predicted by partisan theory. In this case, the empirical evidence points toward higher interest rates under left-wing governments. This effect, however, only appears at higher levels of central bank independence.

This paper also introduces heterogeneous estimators of the forward-looking monetary rule as an alternative to the pooled methods. As central banks are forward-looking, the approach employed in previous studies, that is, to use (ex post) backward-looking data, is not realistic. Under the assumption of forward-looking behavior, the results suggest heterogeneity in the long-run coefficients and only the mean group estimator should therefore be considered. In this case, we do not find a partisan effect and a traditional Taylor rule incorporating forecasts of inflation and the output gap prevails. In contrast to our estimates of backward-looking Taylor rules, our estimates for forward-looking Taylor rules adhere to the Taylor principle.

In a nutshell, our results suggest that the empirical finding in previous studies that interest rates are lower under left-wing administrations is caused by using the same and unrealistic (i.e., backward-looking) Taylor model in panels of countries. When country heterogeneity and a forward-looking central bank behavior are considered, there is no clear evidence in support of a partisan effect on monetary policy. Using our preferred models, we also tested the hypothesis that independent central banks assign more importance to inflation only when left-wing governments are in office and elections are near. Our results do not lend support to this theory.

We also do not find support for the view put forward by Clark and Arel-Bundock (2013) that central banks are conditional inflation hawks, that is, they care about inflation, but only when the government is left-wing and in the close vicinity of an election date. There are several potential explanations for why we do not confirm the findings of Clark and Arel-Bundock (2013). First, we are not focusing the analysis solely on the case of the United States, but on a group of more diverse developed countries in which the convenient separation between "Democrats" and "Republicans" is not possible anymore. Second, our estimation methodology is different and a bit more complex than the ordinary least squares method used in that paper. We use a mean group estimator that allows for slope heterogeneity (Pesaran & Smith, 1995; Pesaran et al.,1999). This estimator also controls for potential biases in the dynamic specification of the model. Third, we consider real-time forward-looking information of both inflation and output gap in our model. Finally, the period of data covered in our analysis is not the same as in Clark and Arel-Bundock (2013). Their analysis starts in the 1950s and covers periods of time in which some of the main variables experienced higher volatility than in more recent times.

Despite its innovations, our study has also limitations. The most important one is that we follow previous studies by assuming that central bank independence is exogenous. In the short term, this seems a reasonable assumption but in the longer term government ideology may affect the level of central bank independence, be it via a traditional partisan effect (right-wing governments favor CBI) or a credibility enhancing effect (left-wing governments favor CBI). We leave this issue for future research.

Furthermore, we have not considered the role of fiscal policy. Theoretically the stance of fiscal policy may be driven by ideological factors and the central bank may respond to the fiscal stance. However, empirical evidence does not clearly suggest that the stance of fiscal policy is driven by partisan factors. In his survey of partisan influence, Potrafke (2017, p. 739) concludes: "Empirical evidence on deficit spending is ambiguous and does not clearly indicate that left-wing governments increased public debt and budget deficits."

In addition, our sample period ends in 2005 as we wanted our study to be comparable with previous research. Future work could expand the sample period, by combining money market rates and shadow rates. The latter is needed to take the effective lower bound into account. As pointed out by Claus, Claus, and Krippner (2016), the shadow short rate is a synthetic summary measure that is derived from yield curve data and essentially reflects the degree to which intermediate and longer maturity interest rates are lower than would be expected if a zero policy rate prevailed in the absence of unconventional policy measures. This measure is better at capturing the stance of monetary policy, especially in the effective lower bound (ELB) period when several central banks started their quantitative easing programs. The shadow rate incorporates this information.

Another important issue for future research is the possibility that central bank behavior may be asymmetric (see, for instance, Brüggemann & Riedel, 2011). Central bank reaction functions (and the reaction coefficients contained therein) can, for instance, be different in periods of expansionary and contractionary monetary policy (see, for instance, Beckmann, Belke, & Dreger, 2016). However, it is important that in examining these issues the reaction functions of central banks are estimated using real-time and forward-looking data.

Also, traditional Taylor rule models can be adjusted to account for international spillovers (see, for instance, Beckmann et al., 2016). According to Taylor (2013), central banks no longer decide on policy rates in an independent way but policy reactions have been increasingly affected by the international environment since then. It would be interesting to examine whether accounting for international spillovers has affected the extent to which political factors influence central bank policies, notably so for a more recent sample period than the one used in the present study.

Finally, as there is some evidence that party affiliation affects central bank governors' policies (Neuenkirch & Neumaier, 2015), it would be interesting to examine this issue for the central banks in our sample, again using forward-looking Taylor rule models.

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SUPPORTING INFORMATION

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APPENDIX A
Descriptive statistics

Table A1 presents summary statistics of the variables used for the backward-looking analysis ranging from 1980Q4 to 2005Q4. Table A2 shows summary statistics of the variables used in the forward-looking analysis for the period starting in 1989Q4. Table A3 shows variation in CBD.

Table A1 Descriptive statistics 1980–2005[a]

Variable	Observations	Mean	*SD*	Min	Max
Interest rate[b]	1,964	9.0	5.2	0.02	37.7
Inflation[b]	2,064	5.7	7.5	−1.4	95.4
GDP gap[c]	2,020	−0.1	1.5	−7.1	6.8
Ideology[d]	2,064	2.9	0.9	1.0	4.0
Central Bank Dependence[e]	1,984	0.5	0.2	0.06	0.81

Notes. [a] Sample of 23 countries for which backward-looking information is available.[b] Annual percentage rate. [c] As percentage of trend output. [d] Scale 1–5, representing ideologies from right wing to left wing. [e] Scale 0–1, represents the inverse of CBI.

Table A2 Descriptive statistics 1989–2005[a]

Variable	Observations	Mean	*SD*	Min	Max
Interest rate[b]	1,060	6.7	4.3	0.02	24.0
Inflation[b]	1,065	2.9	2.8	−1.4	23.4
GDP gap[c]	1,060	−0.1	1.3	−7.0	6.8
Expected inflation 1-year[b]	961	2.9	1.9	−1.1	12.3
Expected GDP gap 1-year[b]	961	−0.3	1.0	−3.3	2.2

(Cotinues)

Table A2 (Continued)

Variable	Observations	Mean	*SD*	Min	Max
Ideology[d]	1,065	3.0	0.9	1.0	4.0
Central Bank Dependence[e]	1,065	0.4	0.2	0.06	0.81

Notes. [a] Sample of 21 countries for which forward-looking information is available.[b] Annual percentage rate. [c] Percentage of trend output. [d] Scale 1–5, representing ideologies from right wing to left wing. [e] Scale 0–1, represents the inverse of CBI.

Table A3 Variation in central bank dependence

	Mean	*SD*	Min	Max.
Overall	0.373	0.235	0.06	0.810
Between	—	0.137	0.143	0.641
Within	—	0.194	−0.018	0.794

Note. Total of 1,365 observations with 21 cross sections and 65 periods.

APPENDIX B

Poolability test

Table B1 presents the results of the poolability test for dynamic panels suggested by Baltagi (2008) and Baltagi and Griffin (1997) for the samples used in Table 1. The null hypothesis is that of homogeneity of the slope coefficients. The results in Table B1 indicate that for the full sample of 23 countries used to estimate a backward-looking Taylor rule the assumption of homogeneity does not hold. For the other samples used in Table 1, the hypothesis of homogeneity cannot be rejected.

Table B1 Poolability tests. Null of homogeneitya

	Backward-looking I	Backward-looking II	Backward-looking III	Forward-looking
F-value	1.82	0.62	1.18	0.7
Critical *F* at 1%	1.32	1.39	1.43	1.36
Outcome	Rejected	Not rejected	Not rejected	Not rejected
Countries	23	16	15	21
Observations [b]	1,841	1,340	791	895

Notes. [a] Implemented mainly through the *ivregress* command in STATA. These statistics may be affected by the instability of the individual country regressions. The *F*-statistic is based on the absolute value of the difference between the sum of squares of the restricted and the unrestricted models as stated in section 3. [b] Due to the particularities of the different tests used, the final number of observations in each column might not necessarily be the same as its corresponding results in Table 1.

APPENDIX C

Sensitivity analysis

As an additional sensitivity check, we have dropped each country from the sample in the backward-looking analysis as reported in column (1) in Table 1. Table C1 presents some of the results. It is clear from Table C1 that when countries such as Iceland, Greece, Portugal, or Spain are removed, the size and statistical significance of the slope coefficients change, confirming our conclusion that the sample is not homogeneous.

Table C1 The impact of government ideology on monetary policy without selected countries

	Backward-looking I	No ISL	No GRC	No PRT	No ESP
Ideology	−0.018	−0.005	0.004	−0.015	−0.020
Central Bank Dependence	0.164***	0.170***	0.152***	0.182***	0.179***
CBD*Ideology	−0.063**	−0.072**	−0.048*	−0.069	−0.058*
Output gap	0.122***	0.109***	0.129***	0110***	0.103***
Inflation	0.111***	0.131***	0.322***	0.249***	0.260***
Lagged Dep. Var.	0.786***	0.892***	0.806***	0.822***	0.831***
Number of *N*	23	22	22	22	22
Observations	1,884	1,810	1,801	1,809	1,808

Notes. See notes to Table 1. Iceland, ISL; Greece, GRC; Portugal, PRT; Spain, ESP; Italy, ITA.

More importantly, the marginal effect of ideology on the monetary policy variable loses relevance when some specific countries are excluded. Table C2 presents a clear example of this fact. After excluding Greece, government ideology no longer affects the short-term interest rate, regardless of the value of CBD (cf. Belke & Potrafke, 2012).

Table C2 Marginal effects of ideology on interest rate excluding Greece (conditional on CDB)

Minimum CBD	0.107* (0.066)
Average CBD	0.005 (0.252)
Maximum CBD	−0.065 (0.066)

Note. See notes to Table 1. Standard errors in parentheses.

These results show that the presence of a traditional partisan effect in the monetary policy is sensitive to the inclusion (exclusion) of specific countries into (out of) the analysis, which can be a signal of heterogeneity in the slope coefficients.

APPENDIX D

Adjustments to the sample size

In order to reassure that the conclusions drawn from comparing the results in columns 3 and 4 of Table 2 are not driven by different sample sizes, we have conducted additional adjustments in order to obtain the exact same number of observations in both models. This was only possible at the expense of reducing the available number of observations and cross sections.29 As Table D1 shows, using these samples yields similar results as reported in Table 2.

Table D1 The impact of government ideology on monetary policy. Equal sample size

	Backward-looking (homogeneous)	Forward-looking (homogeneous)
Ideology	0.051*	0.065**
Central Bank Dependence	0.149***	0.145***
CBD*Ideology	−0.088***	−0.080***
(Expected) Output gap	0.111***	0.047

(Continues)

[29] Australia is excluded from the samples.

Table D1 (Continued)

	Backward-looking (homogeneous)	Forward-looking (homogeneous)
(Expected) Inflation	0.011	0.061
Lagged Dep. Var.	0.890***	0.895***
Number of countries	14	14
Observations	723	723
Marginal effects		
Minimum CBD	0.195*** (0.058)	0.196*** (0.060)
Average CBD	0.052* (0.028)	0.065** (0.029)
Maximum CBD	−0.103* (0.056)	−0.074 (0.055)

Note. See notes to Tables 1 and 2.

APPENDIX E

Mean Group average long-run effect

In the context of the MG model proposed in Pesaran and Smith (1995) and Pesaran et al. (1999), λ_i and β_i are varying across groups according to the specification given in Equation 4:

$$\lambda_i = \lambda + \xi_{1i}\, , \beta_i = \beta + \xi_{2i},$$

where ξ_{1i} and ξ_{2i} are assumed to have zero means and constant covariances. This is a standard formulation of the RCM and it introduces heterogeneity through the 'short-term' coefficients $\lambda_i$and β_i. If the long-run effects, $\theta_i = \beta_i / (1 - \lambda_i)$, and the mean lags, $\psi_i = \lambda_i / (1 - \lambda_i)$, also vary randomly across groups with zero mean and constant covariances, then:

$$\bar{\beta} = \frac{1}{N} \sum\nolimits_{i=1}^{N} \beta_i, \quad \bar{\lambda} = \frac{1}{N} \sum\nolimits_{i=1}^{N} \lambda_i, \quad \bar{\theta} = \frac{1}{N} \sum\nolimits_{i=1}^{N} \theta_i.$$

In other words, the 'average' long-run effects of the regressors can now be defined either in terms of the 'average' of the short-run slope coefficients, $\bar{\beta} / \left(1 - \bar{\lambda}\right)$, or the average of the long-run ones, $\bar{\theta}$ (Pesaran & Smith, 1995). If $N \to \infty$, the two estimators converge and $\bar{\theta}$ becomes the appropriate measure of the average long-run effect.

Table E1 presents the complete results for the regressions shown in Table 3.

Table E1 Regression results. Heterogeneous estimators. Forward-looking central bank

	PMG		MG	
	Coefficient	*z*-Value	Coefficient	*z*-Value
Long run				
Ideology	0.893***	4.04	−0.396	−0.4
Central Bank Dependence	0.165	0.41	0.184	0.7
CBD*Ideology	−0.535*	−1.87	−2.517	−1.20
Output gap	3.223***	5.29	1.617*	1.88
Inflation	2.628***	11.2	1.872**	2.24
Error correction speed	−0.090***	−6.59	−0.214***	−5.55
Short run				

(Continues)

Table E1 (Continued)

	PMG		MG	
	Coefficient	***z*-Value**	**Coefficient**	***z*-Value**
D. Ideology	−0.024	−0.09	0.293	1.3
D. Central Bank Depend.	0.045	0.82	−0.037	−0.45
D. CBD*Ideology	0.485*	1.7	0.574*	1.79
D. Output gap	−0.011	−0.09	−0.012	−0.07
D. Inflation	−0.043	−0.34	−0.026	−0.21
Number of *N*	21		21	
Observations	935		935	

Notes. Long-term dependent variable: short-term interest rate. Ideology and central bank dependence interacted (normalized). Statistical significance: *significant at 10%; **at 5%; ***at 1%.

APPENDIX F

Mean Group average long-run estimators for testing Clark and Arel-Bundock's hypothesis.

Table F1 presents the regressions used to obtain the marginal effects shown in Table 5.

Table F1 The role of elections. Mean group estimators with a forward-looking central bank

	Low CBD		High CBD	
	Coefficient	***z*-Value**	**Coefficient**	***z*-Value**
Long-run				
Ideology	−1.673	−1.44	−0.086	−0.06
(Expected) Output gap	2.828***	4.32	1.773**	2.44
(Expected) Inflation	1.946**	2.25	1.319	0.71
Before_elec	−1.477	−0.96	−3.447**	−2.27
Ideology*Before_elec	1.807*	1.86	−0.439	−0.33
Error correction speed	−0.174***	−3.00	−0.232***	−4.40
Number of *N*	11		10	
Observations	503		432	

Notes. Long-term dependent variable: short-term interest rate. Ideology and expected inflation interacted (normalized). Inflation and Output gap figures are forward looking. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

Table F2 presents the results of running an alternative specification of our model with a triple interaction term between government ideology, CBD, and the time of elections. These figures are the base for computing the marginal effects shown in Table 6.

Table F2 Alternative specification with triple interaction. Mean group estimators with forward-looking central bank

	Coefficient	*z*-Value
Long-run		
Ideology	1.112	0.38
Central Bank dependence (CBD)	0.448**	2.13
Ideology*CBD	−6.189	−1.19
Expected inflation	1.071	0.76
Output gap	0.871	0.73
Before_elec	−4.470**	−2.05
CBD*Before_elec	0.115	0.28
Ideology*Before_elec	0.941	0.43
Ideology*Before_elec*CBD	2.518	1.57
Error correction speed	−0.257***	−5.52
Number of *N*	21	
Observations	935	

Notes. Long-term dependent variable: short-term interest rate. Ideology and expected inflation interacted (normalized). Inflation and Output gap figures are forward looking. Statistical significance: *significant at 10%; **at 5%; ***at 1%.

APPENDIX G

Variable "*Before_elec*" for testing Clark and Arel-Bundock's hypothesis

The "Before_elec" variable is a dummy variable. For each country, the historical average number of periods between elections is calculated. This figure is divided then by two and the outcome will be called "half-term", which is expressed in quarters (see Table G1). The variable "Before_elec" equals 1 in those cases for which after an election the number of periods exceeds the half-term. This is in order to avoid the uncertainty attached to the election date in many parliamentary systems. For countries in which the president is the major political figure or where no prime minister is elected, such as France, Portugal, and the United States, only the presidential elections are considered.[30] Election dates come from the election reports of the Inter-Parliamentary Union (IPU).

Table G1 Summary of elections—Period 1989Q4 to 2005Q4

	# elections	# quarters	Average half-term (quarters)
AUS	6	65	5
AUT	5	65	6
BEL	4	65	8
CAN	4	65	8
CHE	4	65	8
DEU	5	65	6
DNK	5	65	6
ESP	5	65	6
FIN	4	65	8
FRA*	2	65	16
GBR	4	65	8

(Continues)

[30]Considering parliamentary elections for these particular countries instead does change neither the results nor the conclusions. Results are available on request.

Table G1 (Continued)

	# elections	# quarters	Average half-term (quarters)
GRC	6	65	5
IRL	3	65	10
ITA	4	65	8
JPN	6	65	5
NLD	4	65	8
NOR	4	65	8
NZL	6	65	5
PRT*	3	65	10
SWE	4	65	8
USA*	4	65	8

Notes. Average half-term calculation does not include the election period. Average half-term figures have been rounded.* Presidential elections considered.