Monetary Policy Decisions by the World's Central Banks using Real-Time Data

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Abstract
Do Central Bank (CB) actions follow their words? To address this question, this paper contributes to the literature on the conduct of monetary policy in five directions. First, it derives an optimal monetary policy rule determined by both backward and forward-looking macroeconomic variables, based on a two-equation model for inflation and activity, and a CB loss function. This extended monetary policy rule embodies CB decisions based on both past realizations and forecasts of future values. Second, the empirical specification for the policy rule is a direct application of the analytical expression derived in the theoretical model. Third, we use real-time data available to CBs at the time they take their policy decisions. Fourth, we estimate the model using an unbalanced world panel of monthly 1990-2021 data for 29 countries. Finally, we apply a battery of different econometric specifications and estimation techniques and perform nested tests for identifying differences in the conduct of monetary policy for different time periods and country groups. The empirical results are very supportive of the nested model for the reaction of CBs to real-time information on past realizations and forecasts of inflation and activity variables.

Optimal monetary policy, Taylor rules, heterogeneous panels

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1. Introduction

Econometric assessment of monetary policy rules, such as the Taylor rule, has become an important field of study in the modern monetary policy literature. Optimal monetary policy is predicated on forward-looking actions based on real-time forecasts for future variables. This approach is often reflected in central bank (CB) policy statements and documents.[3]

However, monetary policy actions could also respond to real-time data on current (or past) variables. Uncertainty about forecast accuracy and the predictive power of current and past information for future data realizations could lead CBs to use backward-looking data in taking policy decisions. In addition, actual policy making could simply contradict CB statements committed to forward-looking behavior. Several CBs emphasize both current (i.e., backward-looking) variables and forecasts of future variables in setting their policy rates.[4]

So, do deeds follow words? And which CBs' words? This paper tests for the role of real-time information on both past data and forecasts of future variables in setting monetary policy rates in the world.

We contribute to the literature on the conduct of monetary policy in five directions.

First, we derive an analytical framework for an optimal monetary policy rule determined by both backward and forward-looking macroeconomic variables. This nested specification allows to assess if monetary authorities consider forward and/or backward-looking variables when setting interest rates. Our model is a hybrid variant of a simple New Keynesian model based on two seminal papers that model the optimization problem faced by CBs: Svensson (1997) (a purely backward-looking model) and Woodford (2003) (a purely forward-looking model).

We specify a stylized infinite-horizon New Keynesian model for a closed economy that combines backward and forward-looking versions for inflation and activity, based on an aggregate supply equation (or Phillips curve) and an aggregate demand equation (or IS curve), and adding a standard quadratic loss function for the CB. By solving the intertemporal optimization problem, we obtain a closed-form solution for the CB's

[3] Emphasizing the Fed's forward-looking behavior, Ben Bernanke has said that "... it's a true real-time rule – that is, it estimates reaction functions given actual forecasts available at the time the policy decision was taken" (cited in Aso et al., 2010). The Bank of England states: "taken together, all of these impacts on financial markets and associated changes in expectations affect spending decisions and inflationary pressures in the economy" (Bank of England, 2023).

[4] The Reserve Bank of New Zealand states that "the Monetary Policy Statement … has an assessment of current and projected future economic conditions, including inflation and employment" (Reserve Bank of New Zealand, 2023). The CB of Chile states: "monetary policy decisions … are based on current inflation and inflation projections …" (our translation, Central Bank of Chile, 2020).

monetary policy rate, specified as a linear function of four key variables: past inflation, the inflation forecast, past activity, and the activity forecast.

Second, our empirical specification is a direct application of the analytical expression for the monetary policy rule derived in our theoretical model.

Third, we use real-time data available to CBs at the time they take their policy decisions, in order to address the critique, first made by Orphanides (2001), to using data not known by CBs when they take their decisions. This set-up allows identification of the weights placed by CBs on past information and on forecasts about future variables when taking their policy decisions. To our best knowledge, such a nested model has not been tested yet.

Fourth, we estimate our model using an unbalanced world panel of monthly data that extends from 1990 to 2021 and comprises 29 advanced and emerging economies. The latter countries represent 63% of the world's 2021 GDP (at PPP). Our sample comprises a maximum of 7,822 monthly observations. Other panel studies on (non-nested) monetary policy rules include at most 3,800 observations.

Fifth, we apply a battery of different econometric specifications and estimation techniques to our nested monetary policy rule. In order to check for robustness of our results, we allow for different assumptions on the data-generating process represented by different estimation methods. We start by assuming monetary policy homogeneity over time and across countries. First we apply the Instrumental-Variable Fixed-Effects estimator and the Generalized Method of Moments estimator in its dynamic system version. Then we use three alternative estimators for heterogeneous panels: the Mean Group, Pooled Mean Group, and Augmented Mean Group estimators. Then we test for error-correction models that allow for differences between short and long-run coefficients, using three alternative estimators: the Dynamic Fixed Effect, Pooled Mean Group, and Mean Group estimators. To our knowledge, most of the latter estimators have not been used in empirical studies of monetary policy rules yet.

Then we allow for policy heterogeneity, estimating nested specifications of our monetary policy rule that identify possible differences in the conduct of policy over different time periods and across different country groups. Here we address three separate questions: has monetary policy changed since 2002? Has monetary policy changed since the onset of the Global Financial Crisis? Does inflation targeting make a difference for the conduct of monetary policy?

Finally, we check if our estimations of the world's extended monetary policy rules satisfy the Taylor principle, i.e., if CBs contribute to stabilizing the economy by raising interest rates in response to higher inflation by more than one-to-one (Taylor, 1993a, 1993b).

The empirical results are very supportive of the nested model for the reaction of CBs to real-time information on past realizations and forecasts of inflation and activity variables.

This paper is laid out as follows. In section 2 we review the analytical and empirical literature on monetary policy rules that is relevant to this paper. Section 3 derives an extended optimal monetary policy rule. In section 4 we present our empirical specifications, estimation methods, and data, followed by our empirical findings. Section 5 concludes.

2. Literature Review

John B. Taylor shows in his seminal work (Taylor, 1993a, 1993b) that a simple monetary policy rule – the one that subsequently carries his name – fits appropriately the conduct of policy by the US Federal Reserve. The Fed raises its policy interest rate when inflation exceeds a 2% implicit inflation target or when real GDP exceeds potential GDP. Moreover, Taylor shows that monetary policy effectiveness requires CBs to raise the policy rate by more than one-to-one in response to higher inflation. This policy effectiveness requirement has been termed as the Taylor principle.

Taylor's modelling, simulation, and estimations led to a large body of subsequent theoretical and empirical research that has focused on CB behavior and policy rate rules.[5]

Several seminal papers have developed theoretical models to analyze CB behavior, deriving optimal monetary policy rules. One strand of the literature has derived optimal rules as a result of constrained optimization of a (typically ad hoc) CB loss function (e.g., Svensson, 1997; Clarida et al., 1999; Woodford, 2003). A second strand has derived optimal rules as the result of constrained maximization of a representative agent's welfare in a general-equilibrium framework (e.g., Woodford, 2003; Khan et al., 2003; Galí, 2015; Benchimol and Fourcans, 2019).[6] Woodford (2003) shows that a log-linear approximation of the structural equations of second-strand models implies the same restrictions of first-strand models, and that a quadratic approximation of expected utility takes exactly the form of a CB loss function. Therefore the first-order conditions that characterize optimal policy in second-strand models are equivalent to those in first-strand models.

The core specification of empirical versions of first-strand models comprises a CB loss function, an aggregate-supply equation for the rate of inflation (or "New Keynesian Phillips curve") and an aggregate-demand equation for the output gap (or the "intertemporal IS relation"). In this core model the optimal interest rate is determined by a combination of the deviation of inflation from an inflation target and the output gap (Rotemberg and Woodford, 1997; Svensson, 1997; Clarida et al., 1999; Woodford, 2003). In several empirical

[5] Davig and Leeper (2007) generalized the Taylor principle to an environment in which reaction coefficients in the monetary rule are subject to regime changes, confirming the principle's validity for long-term equilibria.

[6] Optimal monetary policy is widely analyzed in the literature using New Keynesian models, which assume that agents' expectations about the future are rational. Blanchard (2018) criticizes this assumption as unrealistic. Recent research has relaxed the rationality assumption in favor of bounded rationality (e.g., Benchimol and Bounader, 2019).

model extensions – including – models in the second strand – additional potential determinants of the policy interest rate are added, including the exchange rate (Corsetti et al., 2010; Käfer, 2014), housing prices (Adam and Woodford, 2012; Käfer, 2014), asset prices (Galí, 2014), and macroprudential variables (Smets, 2014).

Theoretical models of optimal monetary policy can be classified further into two other groups. The first category comprises models where the interest rate responds only to backward-looking macroeconomic variables, consistent with backward-looking aggregate-supply and aggregate-demand relations.[7] The second group comprises models where the interest rate responds only to forward-looking macroeconomic variables, consistent with forward-looking aggregate-supply and aggregate-demand functions).[8] However, to the best of our knowledge, there are no theoretical studies that derive an optimal monetary policy rule where both backward as well as forward-looking specifications lead to a monetary policy rule where interest rates are set in reaction to both backward and forward-looking macroeconomic variables. This stands in contrast to the behavior of CBs that consider both backward and forward-looking macroeconomic conditions when taking monetary policy decisions.

The empirical literature on monetary policy rules is large. Here we review selectively some key issues on empirical rules that are relevant to our paper. Three papers have surveyed the earlier literature. Sims (2001) evaluated empirical Taylor rules according to their micro foundations and how well the models fit the data. Orphanides (2007) reviewed the development and features of Taylor rules in comparison to alternative monetary policy specifications. Muscatelli and Trecroci (2008) surveyed empirical evidence on the role of political-economy variables and monetary regimes (e.g., central bank independence and inflation targeting) in shaping monetary policy rules.

As in the theoretical literature, there are forward-looking rules (econometric specifications based on forecasts or expectations of future policy determinants) and backward-looking rules (econometric specifications based on actual realizations of policy determinants). Some papers compare empirical rules based on backward-looking information to rules based on forward-looking variables. These evaluations, summarized next, are performed for separate policy rules, not for policy rule models that nest both backward and forward-looking variables.

Some papers show that empirical forward-looking rules are preferable to backward-looking rules due to lags in monetary transmission (e.g., Clarida et al., 2000; Orphanides, 2004). This would be consistent with CB statements about their forward-looking reaction to expected future economic conditions. Yet the evidence is mixed. Rudebusch and Svensson (1999) find

[7] A backward-looking structure is presented in Svensson (1997), Clarida et al. (1999), Ball (1999), Dolado et al. (2005), and Sznajderska (2014), among others.
[8] A forward-looking structure is presented in Giannoni and Woodford (2003), Bullard and Mitra (2002), Woodford (2003), and Galí (2015), among others.

that forecast-based rules marginally outperform rules based on contemporaneous (i.e., backward-looking) information. Smets (1998) finds that the evidence on backward-looking rules is marginally better that on forward-looking rules. Taylor (1999) concludes that there is not much difference between the performance of an inflation forecast and past inflation in his policy rule. To the best of our knowledge, there are no papers that report evidence for broader empirical specifications that include both backward and forward-looking macroeconomic conditions.

Regarding the data used in empirical models, monetary policy rules were initially estimated on data that was not available at the time of the corresponding CB policy decision. Orphanides (2001) correctly criticized such methods, which led to several subsequent estimations based on real-time data and forecasts (e.g., Orphanides, 2001; Tchaidze, 2001; Molodtsova et al., 2007; Bems et al., 2021).

Most empirical Taylor rules have been estimated on time-series data applying econometric or VAR models for individual countries, initially for the U.S. and subsequently for a growing number of mostly advanced economies. Most papers use inflation and output gap measures, and add a monetary policy lag. In most papers, estimated parameters exhibit expected signs and are statistically significant (Lubik and Schorfheide, 2007). Some empirical papers have estimated the relative weights of the key macroeconomic determinants and how they vary across different time periods and countries. Levin et al. (2003) investigate the performance of forecast-based monetary policy rules using five macroeconomic models that reflect a wide range of views on aggregate dynamics. They identify that rules respond to the one-year ahead inflation forecast and to the current output gap, considering significant policy inertia. In contrast, rules with longer forecast horizons are less robust in predicting policy decisions.

Divino (2009) uses a cross-country time-series panel model that provides empirical evidence on monetary policy rules for 16 OECD countries during 1979-1998. However, he only uses backward-looking variables and does not take into account the Orphanides critique.[9] To the best of our knowledge there are no cross-country panel data studies of empirical monetary policy rules based on a specification that nests backward and forward-looking policy determinants and/or use only real-time data in their empirical implementation.

3. An Extended Monetary Policy Rule

In this section, we present an analytical framework where both forward-looking and backward-looking variables are included in the model. This nested specification allows to assess if monetary authorities consider forward and/or backward-looking variables when setting interest rates.

[9] There are a few cross-section studies for monetary policy rules. There is one paper that estimates a Taylor rule using a panel for U.S. states (De la Horra et al., 2021) and another that estimates a Taylor rule using regional averages (Hofmann and Bogdanova, 2012).

We specify a stylized infinite-horizon New Keynesian model for a closed economy, presented in discrete time. Our hybrid variant of a simple New Keynesian model is a combination of Svensson (1997) (a purely backward-looking model) and Woodford (2003) (a purely forward-looking model).[10]

Equations (1) and (2) represent the economy's aggregate supply equation (or Phillips curve) and aggregate demand equation (or IS curve), respectively. Both functions include backward and forward-looking terms for inflation and the output gap:

(1)
$$\pi_t = \alpha_1 E_t \pi_{t+1} + \alpha_2 \pi_{t-1} + \alpha_3 E_t y_{t+1} + \alpha_4 y_t + e_t$$

(2)
$$y_t = E_t y_{t+1} + \beta_1 y_{t-1} - \beta_2 (i_t - E_t \pi_{t+1} - \bar{r}) + \omega_t$$

where π_t is the inflation rate in period t, y_t is the output gap defined as the (log) output relative to the (log) trend output, i_t is the monetary policy rate in period t, $\bar{r}$ is the long run real interest rate, E_t is the expectation operator given information at period t, and $E_t \pi_{t+1}$ is the one-period ahead private-sector inflation expectation. The real interest rate, r_t, is defined by the Fisher equation: $r_t = i_t - E_t \pi_{t+1}$, which converges to $\bar{r}$ in the long run. e_t and ω_t represents supply and demand and shocks, respectively, which are i.i.d. with mean 0 and positive variance. The coefficients α_k and β_k, for k ∈ {1, …, 4}, are positive.

We specify the following standard quadratic loss function for the CB:

(3)
$$L_t{}^{CB} = \frac{1}{2}((\pi_t - \bar{\pi})^2 + \lambda_1 (y_t)^2 + \lambda_2 (i_t - i_{t-1})^2)$$

where $\bar{\pi}$ denotes the inflation target level (or the average inflation level for countries that do not announce an explicit inflation target). The CB preference coefficients λ_k, for k ∈ {1, 2}, are positive and reflect the relative weights or costs (relative to the cost of the deviation of output from its trend level) of the deviation of inflation from average inflation or the inflation target (Svensson, 1997) and of the current deviation of the CB's monetary policy rate from its previous level (Woodford, 2003).[11]

The CB chooses in period t a sequence of inflation rates, output, and monetary policy rates that minimize the expected discounted value of its period loss functions:

(4)
$$E_t \sum_{\tau=t}^{\infty} \beta^{\tau-t} L_t{}^{CB} (\pi_\tau, y_\tau, i_\tau)$$

[10] Recently, Galí (2015) and Nakata and Schmidt (2022) propose a model with an empirical specification that is similar to ours. However, they do not solve and derive an optimal monetary policy rule for a specification that considers both forward and backward variables while setting interest rates.

[11] A more detailed discussion of CB loss functions and applications can be found in Bems et al. (2021).

where β is the CB's period discount rate.

In order to solve equation (4), subject to equations (1) and (2), we propose the following Bellman equation (5), where V_t denotes the value function in period t:

(5)
$$V_t(\pi_{t-1}, y_{t-1}, i_{t-1}) = \min_{\{\pi_t, y_t, i_t\}} \frac{1}{2}((\pi_t - \bar{\pi})^2 + \lambda_1(y_t)^2 + \lambda_2(i_t - i_{t-1})^2) + \beta\, V_{t+1}(\pi_t, y_t, i_t)$$

We show the detail of the derivation of our results in the Appendix. The first-order conditions of (5) with respect to π_t, y_t, and i_t, are, respectively:

(6)
$$(\pi_t - \bar{\pi}) + \beta V_{t+1,\pi_t}{}'(\pi_t, y_t, i_t) = 0$$

(7)
$$\lambda_1 y_t + \beta\left(V_{t+1,y_t}{}'(\pi_t, y_t, i_t) + \alpha_4 V_{t+1,\pi_t}{}'(\pi_t, y_t, i_t)\right) = 0$$

(8)
$$\lambda_2(i_t - i_{t-1}) + \beta\left(V_{t+1,i_t}{}'(\pi_t, y_t, i_t) - \beta_2 V_{t+1,y_t}{}'(\pi_t, y_t, i_t) - \alpha_4\beta_2 V_{t+1,\pi_t}{}'(\pi_t, y_t, i_t)\right) = 0$$

where $V_{t+1,x_t}{}'(\pi_t, y_t, i_t)$ denotes the partial derivative of V_{t+1} with respect to any variable x_t. In order to solve the first-order conditions we apply a guess-and-verify method. As the loss function has only quadratic terms, the guess is a linear form of the partial derivative of the value function.

Then, using equations (6) - (8) and combining with equations (1) and (2), yields the following sequence of monetary policy rates:

(9)
$$i_t = \bar{r} + \emptyset_1 E_t y_{t+1} + \emptyset_2 y_{t-1} + \emptyset_3 E_t \pi_{t+1} + \emptyset_4 \pi_{t-1} + \emptyset_5 i_{t-1} + \emptyset_6$$

where $\emptyset_k$, for k ∈ {1, . . ., 5}, are positive constants that are non-linear functions (defined in the Appendix) of structural model parameters.

Note that equation (9) satisfies the Taylor principle – the requirement that the CB rises its policy rate by more than the increase in past/expected inflation in order to raise the real ex-post/ex-ante real policy rate – in the short-run if and only if $\emptyset_3 + \emptyset_4 > 1$, which depends on structural model parameters, as defined in the Appendix. Note also that equation (9) satisfies the Taylor principle in the long run if and only if $\frac{\emptyset_3+\emptyset_4}{1-\emptyset_5} > 1$.

4. Empirical Specifications, Methods, and Data

4.1 Empirical Specifications and Estimation Methods

Our empirical specifications are straightforward versions of the analytical expression for the monetary policy rule derived above, applied to a sample of aggregate cross-country time-series panel data observations. We start with the empirical version of equation (9) and then turn to its error-correction version. In order to check for robustness of our subsequent results, we will estimate each specification using different estimation methods.

The following empirical version of equation (9) reflects our extended monetary policy rule that nests backward and forward-looking measures of inflation and activity, and the lagged dependent variable:

$$i_{j,t} = \delta_0 i_{j,t-1} + \delta_1 \pi_{j,t} + \delta_2 \pi^f_{j,t} + \delta_3 y_{j,t} + \delta_4 y^f_{j,t} + \eta_j + \theta_t + \varepsilon_{j,t} \tag{10}$$

where $i_{j,t}$ is the monetary policy rate at time t in country j, $\pi_{j,t}$ is a past inflation measure available at time t in country j, $\pi^f_{j,t}$ is an inflation forecast measure available at time t in country j, $y_{j,t}$ is a past activity measure available at time t in country j, $y^f_{j,t}$ is an activity forecast measure available at time t in country j, η_j is a country-specific factor (which includes variables such as the inflation target or the long term real interest rate), θ_t is a time-specific factor (which captures shocks that are common to all countries), and $\varepsilon_{j,t}$ is a country and time-specific error term. Consistent with the stochastic terms of equation (10), our estimators correct for country unobservable variables (that are constant over time) and time effects (that are fixed across countries).

First we estimate equation (10) by assuming parameter homogeneity across countries, and subsequently allowing for parameter heterogeneity.

We start by using the Instrumental-Variable Fixed-Effects (IV-FE) estimator, which corrects for possible endogeneity in the smoothing term, due to possible autocorrelation in the error term. We use two lags of all independent variables as instruments. Our second estimation technique is the Generalized Method of Moments (GMM) estimator in its dynamic system version (DSGMM), which performs jointly regressions in levels and in first differences, using the Windmeijer variance correction.[12] All independent variables are treated as potentially endogenous, using their first and second lags as internal instruments. We also perform two specification tests: the Hansen test for the null hypothesis of overall validity of instruments and the Arellano and Bond test for first and second-order serial correlation of errors.

Then we relax the assumption of parameter homogeneity, applying regression models that estimate individual country coefficients separately and then report average country coefficients (e.g., Eberhardt and Teal, 2012; Eberhardt et al., 2013). Since our sample is comprised by 29 countries with an average of 270 observations per country, we can fully

[12] Arellano and Bond (1991) propose a GMM estimator used to estimate dynamic models of panel data and present procedures and specification tests for consistent estimation of parameters and their asymptotic covariance matrix for the dynamic system extension of the GMM panel data model. Windmeijer (2005) proposes a finite-sample correction for the variance of the linear efficient two-step GMM estimator, leading to a more accurate inference.

exploit country heterogeneity. We use three estimators. First we use Pesaran and Smith's (1995) Mean Group (MG) estimator, which allows for heterogeneous slope coefficients and reports the average as the coefficient of interest. Then we apply Pesaran's (2006) Pooled Mean Group (PMG) estimator. This method is similar to the MG estimator; however, it includes means of the independent variables as regressors to correct for possible correlations across countries due to common shocks. Our third estimator is Eberhart and Teal's (2012) Augmented Mean Group (AUG-MG) estimator that accounts for cross-section dependence.[13] In using the three estimators we use a linear average of the estimated parameters and, for robustness, a Huber-weighted average that corrects for the presence of outliers.

Now we turn to the second specification, which is the error-correction version of equation (10):

$$\text{(11)} \qquad \Delta i_{j,t} = \varphi + \gamma\left(i_{j,t-1} - \mu_1 \pi_{j,t-1} - \mu_2 \pi^f_{j,t-1} - \mu_3 y_{j,t-1} - \mu_4 y^f_{j,t-1}\right) + \gamma_1 \Delta\pi_{j,t} + \gamma_2 \Delta\pi^f_{j,t} + \gamma_3 \Delta y_{j,t} + \gamma_4 \Delta y^f_{j,t} + \varepsilon_{j,t}$$

where the $\Delta x_{j,t}$ operator for variable $x_{j,t}$ is defined as $\Delta x_{j,t} = x_{j,t} - x_{j,t-1}$, γ is the long-run coefficient, and φ is a constant used in error-correction models. Parameters μ_k for k ∈ {1, . . ., 4} represent the coefficients of the long-run equation for the independent variables, and parameters γ_l for l ∈ {1, . . ., 6} represent the short-run dynamics.

To estimate equation (11) we use three different methods that allow for varying degrees of heterogeneity. First we estimate a Dynamic Fixed Effects (DFE) model, which restricts all parameters to be homogeneous across countries. Then we estimate a Pooled Mean Group (PMG) estimator, proposed by Pesaran et al. (1999), which allows for heterogeneity in short-run dynamics but constrains long-run dynamics to be common to all countries. Finally, we estimate Pesaran and Smith's (1995) Mean Group (MG) estimator, which allows for heterogeneity in all parameters.

4.2 Data

Our data sample is an unbalanced cross-country time-series panel extending from January 1990 to December 2021. The sample includes 29 countries, comprising 13 advanced countries and 16 emerging-market and developing economies; the maximum sample size comprises 7,822 monthly observations. We include 25 inflation-targeting (IT) countries and 4 non-IT countries.[14] Our data frequency is monthly, which is similar to the frequency of

[13] This estimator accounts for the effect of common shocks by the inclusion of a common dynamic process in the country regression.

[14] The 21 countries with IT in place at the end of our sample period adopted this regime in March 2015 (Costa Rica) or before (all other 20 countries). For the IT countries, we only consider the sub-periods since their individual IT adoption and through December 2021, which implies 2,737 monthly observations. Therefore the number of observations of IT countries before they adopted IT and of

monetary policy decisions. The database is assembled with data from national CBs, BIS, OECD, and Consensus Forecast.

We select countries and their corresponding starting month in our sample based on satisfying jointly the following four conditions met by each country or its corresponding CB, met from the starting month and through the last month of our sample period: (i) a modern monetary policy framework, reflected in using a monetary policy rate as the main policy instrument; (ii) one of the following IMF sub-categories of a floating exchange-rate arrangement: independently floating, managed floating, dirty floating, or clean floating; (iii) a sustained or trend inflation rate below 20% per year; (iv) data available for all dependent and independent variables used in the regressions for at least 3 years or 36 continuous monthly observations ending in December 2021. Sample starting dates are reported for each country in Table 1.[15]

Our empirical variables are the following. The dependent variable is the monetary policy rate set by the CB that is observed at the end of the corresponding month t. The backward-looking inflation measure is the annualized three-month Consumer Price Index (CPI) inflation rate ending at t-2 ($Inflation_{t-4,t-2}$). The backward-looking activity measure is the average seasonally-adjusted unemployment rate for the three-month period ending at t-2 ($Unemployment_{t-4,t-2}$).[16] As opposed to the unobservable output gap, this variable is observable and is correlated with estimated measures of the output gap (by Okun's Law).[17]

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the four countries that did not have IT in place at any time in 1990-2021 is 5085. These two sets of observations will be used when testing for differences in policy behavior between IT and non-IT regimes, as reported in Table 7 below.

[15] We exclude countries that do not satisfy the aforementioned conditions. This implies excluding countries like Bolivia and Venezuela, which do not meet requirement (i). All countries with other IMF arrangements, like stabilized arrangements, managed arrangements, all types of exchange rate bands, and conventional pegs, do not satisfy requirement (ii), such as China, Costa Rica, and Egypt. Countries like Argentina, Venezuela, and Zimbabwe are excluded as they do not meet requirement (iii). Uruguay is excluded for not meeting requirement (iv). Finally, for the case of Chile, which replaced an inflation-adjusted MPR by a nominal MPR (like that adopted in all other countries) in June 2001, our starting date is July 2001.

[16] There are eight countries in our sample that do not report the seasonally-adjusted unemployment rate. We use the unadjusted unemployment rate for the latter countries.

[17] To use real-time data, we use the unemployment rate lagged by two periods and more, to make sure that CBs have this information available at the time of their policy decision. For Australia, New Zealand, and Switzerland, only quarterly unemployment rates are published; for the latter we use quarterly information available for the corresponding CB at the date of its policy decision.

Table 1
Sample starting date by countries

Country	Month	Year	Country	Month	Year
Australia	1	1990	Malaysia	1	2006
Brazil	1	2003	Mexico	11	1998
Canada	1	1990	New Zealand	1	1990
Chile	7	2001	Norway	1	1990
Colombia	10	1999	Peru	9	2003
Czech Republic	12	1995	Philippines	1	1999
Eurozone	1	1999	Poland	5	2000
Guatemala	1	2010	Romania	12	2004
Hungary	3	2008	Russia	1	2004
Iceland	4	2001	Sweden	10	1992
India	1	2005	Switzerland	1	1990
Indonesia	6	2005	Turkey	10	2003
Israel	1	2004	UK	10	1992
Japan	1	1990	USA	1	1990
Korea	5	1990			

Note: the full data base, a detailed explanation of data selection and construction procedures, and the regression files are available from the authors by request.

Our forward-looking variables are the CPI inflation forecast and the GDP growth forecast.[18] The horizon for both forecasts extends over a twelve-month period starting with the current month t and ending eleven months into the future, at month t+11.[19] Our choice of one-year-ahead forecasts is consistent with CB practice and research findings (such as Levin et al., 2003, quoted above, and Taylor, 1993a, 1993b) for forward-looking monetary policy rules, which report that results based on one-year ahead inflation forecasts are more robust than rules with longer forecast horizons.

[18] Other activity forecasts – for the unobservable output gap or the observable unemployment rate – are not available for our full sample.

[19] CBs use different categories of real-time information (available at the time of CB policy decisions) on future expectations and forecasts about inflation and activity variables, including market-based implicit inflation expectations (derived from spreads between indexed and non-indexed Treasury or CB bonds at different maturities), survey-based forecasts (like Consensus Forecast data or data from national surveys, including those conducted by CBs), and internal CB forecasts. For reasons of data available to all CBs and consistency of data over time and across countries, we use forecasts from Consensus Forecast.

Our forecast measures are based on Consensus Forecast data. Since the latter are published for calendar years, we construct a weighted average of two calendar-year forecasts, relevant for month m.

Our monthly one-year ahead inflation forecast is defined as:

(12)
$$\pi_{j,t}^{f} = \frac{(12-\mathrm{m}+1)}{12} * \text{CPI inflation forecast for year T} + \frac{(\mathrm{m}-1)}{12} * \text{CPI inflation forecast for year T} + 1$$

where m is the corresponding month, period t is a month, extending over the sample period covering from 1990.1 (t=1) through 2021.12 (t=372), and T is a calendar year covering from 1990 through 2021.

A similar procedure is applied to calculate the GDP growth forecast between t and t+11. Our monthly one-year ahead GDP growth forecast is defined as:

(13)
$$y_{j,t}^{f} = \frac{(12-\mathrm{m}+1)}{12} * \text{GDP growth forecast for year T} + \frac{(\mathrm{m}-1)}{12} * \text{GDP growth forecast for year T} + 1$$

Table 2 reports summary statistics for our full unbalanced sample. The sample average monetary policy rate is 5.1%, its standard deviation is 4.0%, and the range between maximum and minimum monthly observations reflect the largest difference in monetary policy stance across countries and over time. The annualized three-month inflation rate average is 3.7% and the wide range between extreme points reflects the influence of exceptional idiosyncratic inflation and deflation shocks. Average unemployment is 7.2%, the average inflation forecast (3.5%) is lower than actual average inflation, and the average GDP growth forecast is 3.4%.

Table 3 summarizes bi-variate correlations of all variables. Simple correlations between the monetary policy rate and its potential determinants exhibit expected signs, except for unemployment.[20]

Before conducting our empirical analysis, we test for unit roots in our sample. Since we are working with panel time-series data we perform panel data unit root tests.[21] First we apply a Fisher-type test proposed by Choi (2001), which is based on a combination of p-values of the test statistic for a unit root in each cross-section unit (we perform an Augmented Dickey-Fuller test with two lags for each unit). We test for the null hypothesis that time

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[20] As mentioned above, as a result of data availability restrictions we use the unemployment rate as a measure of past activity and the output growth forecast as a measure of activity forecast. This implies that their expected signs in our extended model equations (10) and (11) are negative and positive, correspondingly.

[21] Detailed discussions of panel time-series tests and estimation techniques are in Barbieri (2009) and Smith and Fuertes (2010).

series display a unit root, while the alternative hypothesis is that a fraction of the sample is stationary. We reject the null hypothesis for all our variables.

Table 2
Sample statistics (in percent)
Unbalanced panel of 7,822 monthly observations, 29 countries, 1990.1 - 2021.12

Statistics	Monetary policy rate	Inflation	Unemployment	Inflation forecast	GDP growth forecast
Mean	5.1	3.7	7.2	3.5	3.4
Median	4.7	3.1	7.0	3.0	3.1
St.Dev.	4.0	4.1	4.2	2.8	2.9
Max.	18.1	18.3	16.1	18.9	10.1
Min.	-0.8	-9.9	2.1	-1.2	-7.2

We also apply a test proposed by Pesaran (2007), which corrects for cross-sectional dependence and serially correlated errors; the conclusion is the same as for the previous test. Therefore we do not find evidence of integrated processes in our panel sample. We also test for the presence of common shocks in our data. We perform Pesaran's (2004) test for cross-section dependence in panel time-series data. We reject the null hypothesis of cross-section independence; hence we will consider possible common shocks in some of our subsequent regressions[22].

Table 3
Bi-variate correlations

	Monetary policy rate	Inflation	Unemployment	Inflation forecast	GDP growth forecast
Monetary policy rate	1.00	**0.81**	**0.51**	**0.79**	**0.41**
Inflation	**0.46**	1.00	**0.37**	**0.91**	**0.34**
Unemployment	**0.36**	**0.21**	1.00	**0.39**	0.09
Inflation forecast	**0.72**	**0.50**	**0.41**	1.00	**0.49**
GDP Growth Forecast	**0.30**	**0.22**	0.07	**0.29**	1.00

Notes: (1) upper matrix: cross-section correlations; lower matrix: panel correlations. (2) Correlations statistically significant at 1% are noted in bold.

[22] All our unit root test results are available on request.

5. Empirical Results

We report full results for our two specifications and using different estimators. We start assuming homogeneous policy behavior over time and across space. Then we test for differences in monetary policy behavior in different time periods and across different country groups.

5.1 Results for homogeneous policy behavior

We start by reporting regression results for the extended monetary policy rule of equation (10). First we focus on our results for the IV-FE and DSGMM estimators (Table 4).

Table 4
Panel regressions: Monetary Policy Rate
Specification: equation (10)
Estimation Methods: IV-FE and DSGMM

Dependent variable: Monetary Policy Rate						
Method	IV-FE (1)	IV-FE (2)	DSGMM (3)	DSGMM (4)	IV-FE (5)	DSGMM (6)
Lagged MPR	0.920*** (0.040)	0.911*** (0.039)	0.912*** (0.038)	0.902*** (0.039)	0.907*** (0.038)	0.920*** (0.037)
Inflation	0.021*** (0.003)		0.018*** (0.002)		0.010*** (0.004)	0.009*** (0.002)
Unemployment	-0.013** (0.005)		0.008** (0.006)		-0.016*** (0.006)	-0.010*** (0.006)
Inflation forecast		0.113*** (0.044)		0.113*** (0.024)	0.087*** (0.040)	0.090*** (0.023)
GDP growth forecast		0.026*** (0.007)		0.021** (0.000)	0.010** (0.004)	0.005* (0.001)
Observations	7822	7822	7822	7822	7822	7822
Countries	29	29	29	29	29	29
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Month FE	Yes	Yes	Yes	Yes	Yes	Yes
Specification tests (p values)						
Hansen test	-	-	0.612	0.728	-	0.856

Serial correlation:						
First order	-	-	0.015	0.018	-	0.017
Second order	-	-	0.421	0.441	-	0.411

Notes: for the IV-FE estimator, we instrument the lagged dependent variable with two further lags of the monetary policy rate and of inflation and unemployment. Heteroskedastic and autocorrelated (AR-1) robust standard errors are reported in parenthesis. *** p<0.01, ** p<0.05, and * p<0.1. For DSGMM estimations we report results based on the Windmeijer variance correction. We also perform two specification tests for columns, (3), (4), and (6): the Hansen test for the null hypothesis of overall validity of instruments and the Arellano and Bond test for first and second-order serial correlation of errors. We reject both the Hansen and Arellano and Bond test hypotheses.

In all regressions the lagged dependent variable is highly significant and its coefficient is large, exceeding 0.9 (but significantly smaller than 1) – which is not surprising, considering the sample's monthly frequency. This implies that the size of the long-term effect of a change in any other right-hand side variable is roughly 10 times as large as the corresponding reported coefficient.[23] This strong monetary policy inertia confirms CBs' aversion to changes in their key policy instrument, as reflected in our loss function (equation (3)). Another general result is that country and time fixed effects are significant in all relevant regressions. We will not comment further on the two latter general results, which are confirmed in subsequent regression tables.

Before reporting the results for our extended model in rows 5 and 6, we present results for model versions that are only backward-looking (rows 1 and 3) and only forward-looking (rows 2 and 4). For the backward-looking versions, coefficients display expected signs and are highly significant, except for unemployment in column 3, which exhibits the wrong sign.[24] For the forward-looking versions, all coefficients display expected signs and are highly significant.[25]

However, the previous results are biased due to omission of relevant variables. Hence now we turn to the results for our extended model specification in equation (10), which nests past realizations and forward-looking forecasts (rows 5 and 6). They reflect that all coefficients display expected signs and most are highly significant (at 1% significance levels), except the GDP growth forecast (significant at 10%). The results for the DSGMM estimator (our preferred technique) are similar to those based on the IV-FE estimator.

[23] The long-term coefficients are calculated for the steady state, by multiplying the short-term coefficients by $\frac{1}{1-\gamma}$.

[24] The size of our past inflation coefficient is similar to the coefficient reported for Germany and half the size of the coefficient reported for the U.S. by Molodtsova et al. (2008) for a backward-looking Taylor rule estimated on real-time data.

[25] The size of our inflation forecast coefficient is similar to the coefficient for the U.S. by Boivin (2006) for a forward-looking Taylor rule based on real-time U.S. Greenbook data.

Both past inflation and the inflation forecast are highly significant in all results. A key finding of our extended model results is that the coefficient estimate of the inflation forecast is several times as large as the coefficient of past inflation. In this table the inflation forecast weighs about eight times the size of past inflation. This is validated in most subsequent reported results: the inflation forecast coefficient is about one magnitude (between nine and ten times) larger than the past inflation coefficient. Therefore CBs attach much larger weight to inflation forecasts than to current (i.e., past) inflation when taking monetary policy decisions. This novel result reflects that monetary policy decisions are strongly and significantly forward-looking regarding inflation. The latter result seems consistent with previous work that shows that empirical forward-looking rules are preferable to backward-looking rules (Clarida et al., 2004; Orphanides, 2004).

Current (i.e., past) unemployment is highly significant in determining monetary policy. The GDP growth forecast is significant at a 10% and 5% level. We are not able to assess the relative weights (or coefficients) of our two activity variables because the corresponding measures are not comparable.

Now we turn to results obtained from estimation models for heterogeneous panels, with different variants of the MG estimator based on Huber-weighted averages that correct for the presence of outliers (Table 5). The first row reports MG estimation results, the second shows PMG results, and the third presents AUG-MG results. Note that both the PMG and the AUG-MG estimators control for unobserved common shocks. The results using MC, PMG and AUG-MG methods are similar to those obtained using IV-FE and DSGMM, both in coefficient sizes and significance levels.

Table 5
Panel regressions: Monetary Policy Rate
Specification: equation (10)
Estimation methods: MG, PMG, and AUG-MG

Dependent variable: Monetary Policy Rate			
Method	MG (1)	PMG (2)	AUG-MG (3)
Lagged MPR	0.931*** (0.009)	0.910*** (0.021)	0.903*** (0.029)
Inflation	0.005*** (0.003)	0.011*** (0.002)	0.009*** (0.004)
Unemployment	-0.014** (0.008)	-0.008** (0.005)	-0.014 (0.016)

Inflation forecast	0.014** (0.005)	0.013** (0.005)	0.093*** (0.041)
GDP growth forecast	0.026*** (0.005)	0.008 (0.009)	0.010** (0.004)
Observations	7677	7677	7677
Countries	29	29	29
Country FE	Yes	Yes	Yes
Month FE	Yes	Yes	Yes

Notes: we instrument the lagged monetary policy using two further lags of the monetary policy rate, and of inflation and unemployment. Heteroskedastic and autocorrelated (AR-1) robust standard errors are reported in parenthesis. *** p<0.01, ** p<0.05, and * p<0.1.

We obtain highly significant results for most variables under each MG variant, which are similar to those reported in columns (5) and (6) in Table 4. Only the GDP forecast coefficient estimated by PMG and the unemployment coefficient estimated by AUG-MG are not significantly different from zero. Another idiosyncratic result obtained here, using the first two estimators, is that the size of the inflation forecast coefficient is similar to that of the past inflation coefficient – as opposed to the result reported for the third estimator, for which the inflation forecast estimate is ten times larger than the past inflation coefficient, which is consistent with the estimations results reported in other tables.

Our next results are based on the error-correction (EC) specification consistent with equation (11), in order to discriminate between short-run and long-run policy response coefficients for our extended policy rule. Table 6 shows the results of our EC model, using DFE, PMG, and MG estimators.

Table 6
Panel regressions: Monetary Policy Rate
Specification: equation (11)
Estimation methods: DFE, PMG, and MG

Dependent variable: Monetary Policy Rate			
Method	DFE	PMG	MG
	(1)	(2)	(3)
A. Short run			
Inflation	0.010*** (0.004)	0.012*** (0.001)	0.013*** (0.003)
Unemployment	-0.183***	-0.224***	-0.174***

	(0.005)	(0.004)	(0.006)
Inflation forecast	0.082***	0.189***	0.110***
	(0.004)	(0.005)	(0.003)
GDP growth forecast	-0.008	0.071**	0.062**
	(0.015)	(0.007)	(0.015)
B. Long run			
gamma	-0.053***	-0.032***	-0.063***
	(0.001)	(0.004)	(0.001)
Inflation	0.128***	0.155***	0.120*
	(0.005)	(0.008)	(0.005)
Unemployment	0.063	-0.171***	0.049
	(0.073)	(0.054)	(0.061)
Inflation forecast	1.017***	1.211***	1.246***
	(0.101)	(0.096)	(0.111)
GDP growth forecast	0.661***	0.921***	0.028
	(0.123)	(0.087)	(0.021)
Observations	7793	7793	7793
Countries	29	29	29

Notes: we instrument the lagged monetary policy using two further lags of the monetary policy rate, and of inflation and unemployment. Heteroskedastic and autocorrelated (AR-1) robust standard errors are reported in parenthesis. *** $p<0.01$, ** $p<0.05$, and * $p<0.1$.

The results are consistent with the EC model: the error-correction coefficient gamma is highly significant, 11 out of 12 short-run coefficients are very significant, and 9 out of 12 long-term coefficients are very significant. The size of the gamma coefficient is small – between -0.032 and -0.063 – reflecting that the monetary policy correction of model deviations (or errors) is approximately 5% of estimated deviations. This confirms the significant policy inertia found in monthly data, which is reflected by the large coefficients reported for the lagged dependent variable reported in all other tables for our model consistent with equation (10).

The short-run inflation coefficients are very similar in size to the inflation coefficients reported previously. Moreover, they replicate the previous results on their relative size: the inflation-forecast coefficients are roughly one magnitude larger – between 8 and 16 times – than past-inflation coefficients. The same result about relative size is obtained for the long-run inflation coefficients: inflation-forecast parameter estimates are between 8 and 10 times as large as those for past inflation. Moreover, the absolute size of long-run inflation coefficients is roughly ten times the size of corresponding short-run coefficients. This is consistent – again – with the large and significant policy inertia observed here and in all other reported results.

Most short-term coefficients of activity variables are statistically very significant for the different estimators. For long-term activity variables we obtain three out of six very significant coefficients.

5.2 Results for heterogeneous policy behavior

Our final evidence is on differences in monetary policy behavior observed in different periods and country groups. We address three questions.

(i) As most of our sample CBs adopted a modern policy framework during the 1990s or early 2000s, has monetary policy changed since 2002?[26]

(ii) During the Global Financial Crisis (GFC), monetary policy was actively used to respond to deteriorating financial conditions and the ensuing Great Recession. Since the onset of the GFC and until the end of our sample period (December 2021), monetary policy rates appear to be lower than before the GFC and monetary policy seems to be more dovish than before the GFC.[27] Is this reflected in our results?[28]

(iii) Under inflation targeting (IT), CBs could respond more strongly to inflation than CBs that have other monetary regimes in place. Is this observed in our sample?

To address these issues, we compare monetary policy behavior before and after central banks have adopted a modern policy framework (question i), before and since the GFC (question ii), and between IT and no-IT countries and periods (question iii). We nest the differential policy behavior during particular periods and country experiences within the general specification of our extended monetary policy rule. Our specification is equation (10) for the monetary policy rate, extended here by including interaction terms between

[26] A monetary policy framework comprises CB institutional and operational features that enable and guide a conduct of monetary policy at the frontier of international best practice. Such a monetary policy framework is based on three pillars: (i) CB independence and accountability, (ii) a developed policy and operational strategy, and (iii) policy transparency and communication with markets and the public. In the 1990s or early 2000s, our sample countries adopted a policy framework that is consistent with the abovementioned pillars; advanced economies implemented it earlier and developing economies came somewhat later (Unsal et al., 2022). We select 2002 as the cut-off year, since it marks the year in which all (or most) of our sample countries have an advanced monetary policy framework in place.

[27] Consistent with general practice, we define September 15, 2008 (the day Lehman Brothers collapsed) as the starting date of the Global Financial Crisis. Therefore the dummy variable for the beginning of the GFC is set at 1 from October 2008 to December 2021 (Ramskogler, 2014).

[28] Monetary policy rules are less useful when the monetary policy rate is constrained by the effective lower bound. Then orthodox monetary policy is complemented by the use of heterodox instruments, such as quantitative easing and forward guidance (Bernanke 2015, 2017). In several sample countries, monetary policy rates were held at the effective or zero lower bound during some sub-periods between the GFC and 2021.

each independent variable and a time-period dummy or a country-group dummy (termed "Dummy"). The upper section of Table 7 reports the results that are representative of the full sample. The lower section presents the differential results for the interaction terms between independent variables, on the one hand, and time-period or country-group dummies, on the other. The results are based on the DSGMM estimator and therefore can be compared to the DSGMM results for the non-nested specification, reported in column (6) of Table 4.

Table 7
Panel regressions: Monetary Policy Rate
Specification: equation (10)
Estimation method: DSGMM

Dependent variable: Monetary Policy Rate			
Time periods or country groups	Since 2002 (Jan. 2002-Dec. 2021) (1)	Since GFC (Oct. 2008-Dec. 2021) (2)	25 Inflation-Targeting Countries (3)
Lagged MPR	0.900*** (0.002)	0.951*** (0.002)	0.958*** (0.003)
Inflation	0.009*** (0.003)	0.013*** (0.002)	0.005* (0.003)
Unemployment	-0.011 (0.007)	-0.021** (0.006)	-0.010 (0.007)
Inflation forecast	0.043*** (0.003)	0.045*** (0.002)	0.037** (0.003)
GDP growth forecast	0.007 (0.006)	0.009 (0.005)	0.035*** (0.006)
Dummy	-0.127*** (0.004)	-0.155*** (0.006)	-0.095 (0.012)
Interaction terms with time periods or country groups			
Lagged MPR	0.049*** (0.003)	-0.041*** (0.004)	-0.014*** (0.004)
Inflation	0.004*** (0.001)	-0.008** (0.002)	0.012*** (0.001)
Unemployment	-0.015*** (0.003)	-0.010*** (0.001)	-0.027*** (0.002)
Inflation forecast	0.022** (0.004)	0.026 (0.005)	0.014** (0.006)

GDP growth forecast	0.010*** (0.003)	0.025*** (0.002)	0.009 (0.012)
Observations	7822	7822	7822
Countries	29	29	29
Specification tests (p values)			
Hansen test	0.691	0.891	0.908
Serial correlation:			
First order	0.022	0.026	0.016
Second order	0.383	0.423	0.378

Notes: we instrument the lagged monetary policy using two further lags of the monetary policy rate, and of inflation and unemployment. Heteroskedastic and autocorrelated (AR-1) robust standard errors are reported in parenthesis. *** p<0.01, ** p<0.05, and * p<0.1. We report results with the Windmeijer correction.

The econometric results presented in the upper section of Table 7 for all three rows – which are representative of the full sample – are similar in coefficient signs, sizes, and statistical significance to those reported in column (6) of Table 4. However, some differences arise regarding activity variables: only two of the six coefficients are very significant. This suggests that our extended policy rule does not perform as well for our full data sample, once we consider a specification that nests a differential policy behavior. The latter presumption is confirmed by the results presented in the lower section of Table 7.

The results in the first column of Table 7 show that monetary policy has changed significantly since 2002, compared to the 1990-2001 period. First, the average monetary policy rate is slightly lower – by 12.7 basis points – since 2002. Second, and more important, we find that CBs respond more strongly, significantly, and in the correct direction to all four independent variables in our extended policy rule since 2002. The coefficients of past inflation and the inflation forecast increase by one half, compared to the pre-2002 period. Moreover, in the earlier period CBs did not react to activity variables; since 2002 they respond significantly to past unemployment and the growth forecast.

Now we turn to the differential results found for the period since the start of the GFC, from October 2008 to December 2021, reported in the second column of Table 7. The average monetary policy rate is lower by 15.4 basis points to that prevailing before. CBs react less to past inflation (its coefficient drops by one half) since the GFC, but their reaction to the inflation forecast remains. In contrast to the latter result, CBs react much more strongly to activity variables than before. The coefficient size for unemployment increases by one half and is highly significant. Similarly, the coefficient of the GDP growth forecast – which was not different from zero before the GFC – turns to be large and highly significant since the GFC. These results are a strong confirmation that CBs displayed a significantly more dovish policy since late 2008 and throughout 2021, a period marked by the Great Recession in 2008-09 and the Covid-19 recession in 2021.

We report in the third column the differential results for countries that have IT in place, comparing policy behavior since the country-specific month when CBs adopted IT to the period before IT adoption in the same countries and to CBs that have not had IT in place during the 1990-2021 period.

We find that IT CBs, compared to non-IT banks, exhibit less policy inertia than non-IT CBs: the coefficient of the lagged monetary policy rate is 10% smaller in the former countries. Regarding the core determinants of policy, our results show that IT CBs react more strongly and significantly to three of our four independent variables than non-IT CBs. It is not surprising that IT CBs respond more strongly to both past inflation and the inflation forecast than non-IT CBs. The coefficient of reaction of IT CBs to past inflation (0.016) is highly significant and three times as large as the coefficient of non-IT CBs (0.005). The coefficient of reaction of IT CBs to the inflation forecast (0.044) is also larger than the coefficient of non-IT CBs (0.034). More surprising is our finding that IT CBs also react strongly to the unemployment rate, while the reaction of non-IT CBs to unemployment is not significantly different from zero. Finally, IT CBs share with non-IT CBs the same strong and significant reaction to the GDP growth forecast.

We conclude that our results validate differences in monetary policy behavior observed in different sub-sample periods and between IT and non-IT CBs. First, our extended monetary policy rule is more strongly validated since adoption of a modern monetary policy framework: CBs respond more strongly and significantly to all four independent variables since 2002. Second, CBs exhibit a more dovish monetary policy stance since the GFC than before: they respond less strongly to inflation and more strongly to activity variables since late 2008. Third, IT CBs respond more strongly to past inflation and the inflation forecast than non-IT CBs, and strongly to unemployment as opposed to non-IT CBs. Therefore our extended monetary policy rule is particularly suited for characterizing monetary policy under inflation targeting.

5.3. Taylor Principle

Do our estimated policy rules satisfy the Taylor principle?

Our coefficient estimates for the short-term (monthly) reactions of CBs to a past-inflation shock and to an inflation-forecast shock are significantly smaller than 1. However, CBs exhibit very strong policy inertia in the short term, which is borne by coefficient estimates for the lagged monetary policy rate that are close to 0.90. Both sets of results reflect the fact that CBs respond gradually to an inflation shock. Hence we focus on the long-run response of monetary policy to a sustained change in past inflation and the inflation forecast to address the question about the Taylor principle.

For the level estimations consistent with equation (10), we divide the sum of the coefficients of inflation and the inflation forecast by one minus the coefficient of the lagged dependent variable. Many of our results reported for the case of homogeneous policy behavior satisfy the Taylor principle, implying combined long-term inflation coefficients that are larger than 1. Our preferred result based on the DSGMM estimator (column (6), Table 4) implies a combined long-term inflation coefficient of 1.24, satisfying the Taylor principle. For the results based on the IV-FE estimator (column (5), Table 6), the combined long-term coefficient is 1.04. For heterogenous panels, only the combined long-term inflation coefficient for the AUG-MG results (Table 5, column (3)) exceeds 1. In the case of the error-correction model consistent with equation (11), we focus directly on the sum of the two long-term inflation coefficients (lower section of Table 6). In all three columns, the latter sum is larger than and significantly different from 1.0.

Our last conclusions on the Taylor principle are derived from the results reported for heterogeneous policy behavior (Table 7). While monetary policy before 2002 does not satisfy the Taylor principle, it does so in 2002-2021. This is consistent with the observed increase in inflation and inflation-forecast coefficients, starting in 2002 with the adoption of modern policy frameworks. The opposite is observed when splitting the full sample at the time of the GFC: before the crisis the Taylor principle is satisfied but afterwards – when monetary policy turns more dovish – it is not.

Finally, when splitting the sample between IT and non-IT CBs, we find significant differences between both groups. The sum of long-term inflation coefficients of the policy rule of non-IT CBs is exactly 1.0, and therefore the validity of the Taylor principle is rejected for this group of countries. We noted above that both short-term past-inflation and inflation-forecast coefficients are significantly larger in the group of IT CBs than among non-IT CBs. Hence it is not surprising that the sum of both coefficients is 1.21 in the long term, satisfying the Taylor principle. This suggests that when the explicit objective of monetary policy is on meeting inflation targets, this may enhance policy effectiveness, in comparison to other monetary policy regimes.

6. Conclusions

Optimal monetary policy is predicated on forward-looking actions based on real-time forecasts for future variables. However, monetary policy actions could also respond to real-time data on current and past variables. So, do deeds follow words? And which central bank's words? This paper tests for the role of real-time information on both past data and forecasts about future data in setting monetary policy rates in the world.

We contribute to the literature on the conduct of monetary policy in five novel dimensions. First, we derive an analytical framework for an optimal monetary policy rule determined by both backward and forward-looking macroeconomic variables, based on a two-equation set-up for activity and inflation, and a CB loss function. This extended monetary policy rule nests CB policy decisions based on past realizations and forecasts of future values of inflation and activity.

Second, our specification subject to subsequent empirical testing is a direct application of the analytical expression for the monetary policy rule that is derived theoretically.

Third, we use real-time data available to CBs at the time they take their policy decisions. This set-up allows identification of the weights placed by CBs on past information and on forecasts about future variables when taking their policy decisions.

Fourth, we estimate our model using an unbalanced world panel of monthly observations that extends from 1990 to 2021 and comprises 29 advanced and emerging economies. The latter countries represent 63% of the world's 2021 GDP (at PPP). Our sample comprises a maximum of 7,822 observations.

Fifth, we apply a battery of different econometric specifications and estimation techniques to our monetary policy rule specification that nests backward and forward-looking measures of inflation and activity and the lagged dependent variable. To allow for different assumptions on the data-generating process, we estimate each specification using different estimation methods. First we estimate by assuming policy homogeneity over time and across countries. Then we conduct nested tests for identifying differences in the conduct of monetary policy for different time periods and country groups. Finally we check if our estimations of the world's extended monetary policy rules satisfy the Taylor principle.

The empirical results are very supportive of the nested model for the reaction of CBs to real-time information on past realizations and forecasts of inflation and activity variables. This is confirmed by most regressions for homogeneous and heterogeneous panels, for level and error-correction specifications, and for homogeneous and heterogeneous policy behavior.

In all regressions the lagged dependent variable is highly significant and very large, exceeding 0.9. This implies that the size of the long-term effect of a change in any other right-hand side variable is close to 10 times the size of the reported coefficient. This strong monetary policy inertia confirms CBs' aversion to changes in their key policy instrument, as reflected in our loss function.

Our empirical results for heterogeneous panels (based on MG, PMG and AUG-MG estimators in Table 5)) are consistent with the results for our benchmark regressions for homogeneous panels (based on IV-FE and DSGMM estimators in Table 5). The two latter regressions sets for our level specification are also confirmed by our error-correction specification results (based on DEF, PMG, and MG estimators in Table 6).

Most regression results reflect that the coefficients of past realizations and forward-looking forecasts of inflation and activity variables display expected signs and high levels of statistical significance. We find that both past inflation and the inflation forecast are generally significant and robust determinants of monetary policy decisions. A key finding is that the coefficient estimate of the inflation forecast is several times – on average 10 times – as large as the coefficient of past inflation. This novel result reflects that CBs are strongly forward-looking in their policy reaction to information about inflation: they attach a weight to inflation forecasts that is about a magnitude larger than that attached to past inflation data.

Reflecting limitations on data availability, our activity variables are the past unemployment rate and the GDP growth forecast. Therefore we are not able to compare the relative weights attached by central bankers to the backward and forward-looking activity variables Coefficient estimates of both activity variables generally exhibit expected signs and high significance levels, but they are less robust across different specifications and estimation methods than the coefficients of inflation variables.

Our final evidence is on differences in monetary policy behavior observed in different periods and country groups. First, our extended monetary policy rule is more strongly validated since adoption of a modern monetary policy framework: CBs respond more strongly and significantly to all four independent variables since 2002. Second, CBs exhibit a more dovish monetary policy stance since the GFC than before: they respond less strongly to inflation and more strongly to activity variables since late 2008. Third, IT CBs respond more strongly to past inflation and the inflation forecast than non-IT CBs, and more strongly to unemployment as opposed to non-IT CBs. Therefore our extended monetary policy rule is particularly suited for characterizing monetary policy under inflation targeting.

Finally we turn to the question if our estimated policy rules satisfy the Taylor principle, i.e., if CBs contribute to stabilizing the economy by raising interest rates in response to higher inflation by more than one-to-one. We focus on the long-run response of monetary policy to a sustained change in past inflation and the inflation forecast. Most of our regression results imply long-run inflation coefficients that are larger than 1, validating the Taylor principle.

When considering heterogenous policy behavior across different time periods and country groups, we identify relevant differences. While monetary policy before 2002 did not satisfy the Taylor principle, it does since 2002, when sample countries have in place a modern policy framework. However, the opposite is observed after the Global Financial Crisis, a likely reflection of a more dovish policy stance during the last decade. When splitting the sample between non-IT and IT CBs, only the latter satisfy the Taylor principle – a likely result of enhanced policy effectiveness when the explicit objective of monetary policy is on meeting inflation targets.

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