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## The Effects of Key Parameters of the Monetary Policy Reaction Function on Economic Growth

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# The Effects of Key Parameters of the Monetary Policy Reaction Function on Economic Growth

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**Abstract**

We examine how a more hawkish policy stance – defined as an above-median long-run inflation semi-elasticity of the policy rate – affects economic growth in 37 inflation-targeting countries. To this end, we estimate time-varying, bias-corrected forward-looking Taylor rules for all inflation-targeting countries for which the data permit such estimation. Our results point to sizable growth effects, exceeding 0.5 percent annually, for countries with a more hawkish policy stance. This suggests that the growth benefits reported in the previous literature on inflation targeting are primarily driven by a small subset of countries that react more forcefully to inflation.

**Keywords:** Economic growth, inflation targeting, monetary policy reaction function, hawkishness

**JEL:** E52, E58, O40, O47

# 1 Introduction

Over the past few decades, inflation has remained low and relatively stable in most advanced economies and a growing number of emerging economies. This development is often associated with the widespread adoption of inflation-targeting (IT) regimes, which the empirical literature typically characterizes as following a Taylor rule (see, e.g., Taylor and Davradakis; 2006; Çağlayan and Astar; 2010; Neuenkirch and Tillmann; 2014).[1] By committing to an explicit inflation target and typically granting the central bank sufficient independence to pursue this objective, many countries were able to mitigate the time-inconsistency problem in monetary policy. In the absence of credible rules, policymakers often succumb to the temptation of expansionary monetary policies for short-term gains – such as monetizing fiscal deficits, exploiting short-run Phillips-curve trade-offs, or stimulating investment through lower interest rates. These policies typically lead to rising inflation, which in turn undermines growth, as illustrated most clearly by the stagflation experienced across many advanced economies in the 1970s. Yet, the empirical evidence on the growth effects of IT remains inconclusive. One reason may be that not all inflation-targeting regimes are created equal. For example, El-Shagi and Ma (2023) document that most inflation-targeting countries do not satisfy the Taylor principle: policy rates do not increase more than proportionally with inflation expectations, so real interest rates fail to rise, thereby preventing monetary policy from becoming genuinely contractionary. In this paper, we contribute to this literature by examining how the central bank's commitment to price stability – proxied by the long-run inflation coefficient in an estimated Taylor rule – affects economic growth. Since country fixed effects are a key component of a well-specified empirical growth model, we cannot exploit cross-country variation in the Taylor rule. Instead, we exploit the time variation in monetary policy rules documented in recent work.

We thus contribute to two strands of the literature. First, we add to the empirical literature on the growth effects of monetary policy. While there is a large body of work on how monetary policy regimes – including inflation targeting – shape growth, to the best of our knowledge ours is the first paper to examine how the implementation of a common policy rule in practice affects growth. Second, we extend the literature on time-varying estimates of monetary policy reaction functions (MPRF). A considerable empirical literature has explored time-varying monetary policy reaction functions – typically Taylor rules – for the Federal Reserve, the ECB, and a range of smaller economies. Yet relatively few studies analyze a broad cross-section of countries using a consistent and comparable empirical framework. A notable exception is the study by Anderl and Caporale (2024), who employ a time-varying GMM approach to estimate forward-looking Taylor rules for ten developed economies, using the estimation

[1] The Taylor rule also describes monetary policy behavior reasonably well in economies without formal inflation-targeting frameworks, such as several major Latin American economies (Moura and de Carvalho; 2010) and the United States before an official inflation target was introduced in 2012 (Woodford; 2001; Orphanides; 2003).

strategy introduced by Partouche (2007). Although our ultimate goal is to assess growth effects rather than to analyze time variation in policy rules per se, our sample of 37 inflation-targeting countries makes this the most comprehensive study to date in this line of empirical research.

Our findings indicate sizable and statistically robust effects of a more hawkish policy stance on economic growth, with annual impacts ranging from 0.5 percent to 1 percent across specifications. The growth benefits typically associated with inflation targeting are largely confined to countries that react more forcefully to inflation, implying that adopting an inflation-targeting label alone is insufficient. This underscores the importance of the intensity of policy reactions and suggests that even widely regarded "best-practice" monetary frameworks remain vulnerable to time-inconsistency problems that can undermine growth outcomes.

The remainder of this paper is structured as follows. The next section summarizes the literature. Section 3 introduces our research methodology, and Section 4 discusses the results. Finally, Section 5 concludes.

# 2 Literature review

## 2.1 Monetary policy and growth

The previous literature approaches the impact of monetary policy on growth from two fundamentally distinct perspectives. One strand, the literature on monetary neutrality, examines whether expansionary and contractionary monetary policy have persistent effects on output. The other strand focuses on institutional determinants of growth, analyzing how central bank features such as inflation targeting, central bank independence, and related governance arrangements shape economic performance. Both strands report mixed empirical findings.

Both King and Watson (1997) and Bernanke and Mihov (1998) find that US postwar data is consistent with the neutrality – but not the superneutrality – of money. Yet Mankiw (2001) interprets the very findings of Bernanke and Mihov (1998) as evidence of non-neutrality. More recently, Jordà et al. (2024) find that expansionary policy is neutral, whereas contractions leave permanent scars that reduce output. Serletis and Xu (2024) emphasize that, for most countries, the data does not even permit meaningful tests of neutrality, but they confirm neutrality where such tests are feasible. Serletis and Koustas (1998) test the neutrality and superneutrality of money for ten developed countries and find neutrality in all cases and superneutrality in all but Italy. Italy stands out as the only country in their sample experiencing a persistent shock to money growth. This link between persistent money-growth shocks (and thus persistent changes to inflation) and the failure of superneutrality connects their findings to the broader literature on the growth effects of inflation (Bick; 2010; López-Villavicencio and Mignon; 2011), and, by extension, to the literature on how central bank institutions influence growth through their role in maintaining price stability.

Much of this work focuses on inflation targeting and central bank independence. For inflation targeting, some studies find that it can support long-run growth by anchoring expectations and reducing macroeconomic volatility (Mishkin; 2007; Svensson; 2010), whereas others report little or no effect (Ball and Sheridan; 2004; Brito and Bystedt; 2010; Junankar and Wong; 2020). The evidence on central bank independence is similarly mixed. While some research argues that greater independence strengthens credibility and fosters growth (Akinci et al.; 2015; Garriga; 2016), other studies find no systematic growth benefits (Neyapti; 2001; de Mendonça; 2006; Anastasiou; 2009). A related, largely theoretical literature addresses the more fundamental question of which monetary policy framework – understood broadly as the rule or target variable guiding the conduct of policy – delivers the highest economic growth (see, e.g., Lai and Chin; 2013; Chu and Ji; 2016; Huang et al.; 2017; Chu et al.; 2019; Gil and Iglésias; 2020).

By examining how the operational features of an inflation-targeting framework shape growth, our paper bridges the two literature on policy actions and institutional or rule-based frameworks. Concretely, we ask how a hawkish versus a dovish stance of the central bank within an inflation-targeting framework affects growth. Empirical work on this aspect is still scarce. The paper closest to our question is Gil et al. (2023), who provide theoretical support for the idea that strong policy responses to inflation – which are consistent with the Taylor principle – anchor expectations and create conditions that contribute to growth.

## 2.2 Time-varying monetary policy reaction functions

Our identification strategy exploits time variation in the Taylor rule. This allows us to estimate the effect of changes in Taylor-rule parameter – particularly the long-run inflation coefficient – on growth, while accounting for time-invariant country heterogeneity through fixed effects. Introducing time variation into estimated versions of the original backward-looking Taylor (1993) specification is conceptually straightforward (see, e.g., Judd et al.; 1998; Assenmacher-Wesche; 2006; Kuzin; 2006). However, since the seminal work of Clarida et al. (1998), it has become widely accepted that central banks typically respond to expected inflation and output rather than to current conditions. Future realizations – often used as proxies for expectations under the rational-expectations assumption – are endogenous to monetary policy by construction, since policy is explicitly designed to influence those future outcomes. Orphanides (2001) proposes using real-time forecasts, such as Greenbook or SPF projections, to address this endogeneity concern, since these forecasts are conditioned only on information available at the time and are therefore exogenous to contemporaneous policy shocks. Yet such forecasts are often unavailable, meaning that estimating a forward-looking policy rule requires an instrumental-variables approach that complicates time-varying-parameter estimation.Wesche (2003) proxies inflation expectations using a VAR-based forecast in her Markov-switching framework, and, in a similar spirit, Trecroci and Vassalli (2010) uses trend inflation – derived from a simple filtering procedure – as a measure of long-run expectations in his time-varying-parameter model. Both approaches amount to ad hoc modifications of a standard two-stage IV procedure to accommodate evolving monetary-policy parameters. A two-step procedure for estimating bias-corrected, forward-looking, time-varying Taylor rules was introduced by Kim and Nelson (2006) and has been applied in various empirical settings, including the Brazilian case examined by Carvalho and Muinhos (2023). An alternative approach is the GMM framework proposed by Partouche (2007), which Anderl and Caporale (2024) apply in their ten-country study. A much larger literature examines time variation in monetary policy using VAR-based approaches (see, e.g., Boivin and Giannoni; 2006; Cogley and Sargent; 2005; Primiceri; 2005; Sims and Zha; 2006). However, these approaches do not permit identification of the policy parameters, rendering them unsuitable for addressing our research question.

In this paper, we employ the framework of Kim and Nelson (2006) to model gradual shifts in the reaction function of 37 inflation-targeting countries.

# 3 Model and estimation strategy

## 3.1 Panel growth regression model

**The model** We augment a standard panel growth regression by including a time-varying measure of policy hawkishness as our main explanatory variable.

Following the usual convention in the growth literature, we construct non-overlapping five-year periods and estimate our model on this quinquennial panel (see, e.g., Barro et al.; 2003; Bruns and Ioannidis; 2020):

$$\Delta y_{i,t} = \delta_0 + \delta_1 y_{i,t-1} + \delta_2 \text{Hawk}_{i,t-1} + \phi X_{i,t} + u_i + v_t + \varepsilon_{i,t}, \qquad (1)$$

Here, $y_{i,t-1}$ is the log of real GDP per capita at the onset of the period, and $\text{Hawk}_{i,t-1}$ is an indicator for whether country $i$ exhibited a hawkish policy stance at that time. The vector $X_{i,t}$ contains standard growth determinants, including population growth, investment, government consumption (as a share of GDP), education, and trade openness, all measured at the end of period $t-1$, corresponding to the beginning of the period $t$. Country and time fixed effects are captured by $u_i$ and $v_t$, and $\varepsilon_{i,t}$ is the idiosyncratic error term.

**Estimation** The specification in Equation (1) corresponds to an autoregressive process in the log level of GDP and can be rewritten equivalently as:

$$y_{i,t} = \delta_0 + (\delta_1 + 1) y_{i,t-1} + \delta_2 \text{Hawk}_{i,t-1} + \phi X_{i,t} + u_i + v_t + \varepsilon_{i,t}, \qquad (2)$$

Dynamic panel models such as Equation (2) require estimators that address the endogeneity introduced by the inclusion of the lagged dependent variable. A widely used approach in the growth literature is the system-GMM estimator of Arellano and Bover (1995) and Blundell and Bond (1998), which leverages additional moment conditions to enhance efficiency, especially when the variable of interest exhibits high persistence.

However, the validity of the assumptions required for system GMM when applied to growth regressions has been questioned, particularly the stationarity requirement for the level equation (Moral-Benito; 2013, 2014). Including country and time fixed effects can absorb much of the non-stationarity, as the stochastic trend in output is largely driven by global shocks to total factor productivity. Indeed, the autoregressive coefficient in Equation (2) is well below unity for almost all estimators we consider, indicating that stationarity is not a major concern. Yet to ensure robustness to the Moral-Benito critique, we rely on several estimation strategies rather than a single method.

First, we consider a standard fixed-effects estimator, which merely captures unobserved heterogeneity through fixed effects and leaves other identification issues unaddressed. Second, we use system-GMM to address potential endogeneity. Third, we employ the difference-GMM estimator of Arellano and Bond (1991) as a robustness check against possible violations of the system-GMM moment assumptions. Finally, we use the quasi–maximum-likelihood estimator proposed by Kripfganz (2016), which is asymptotically equivalent to GMM but offers improved finite-sample properties, particularly in the presence of endogenous regressors.

All results are reported in terms of Equation (2), in other words, the tables report $\hat{\delta}_1 + 1$.

**Sample and data** Our sample begins in 2000 and ends in 2020, the close of the most recent five-year block. We include all inflation-targeting countries listed by Central Banks News (2022) for which sufficiently long time series are available. Most of the data for the growth model come from the World Bank's World Development Indicators, with the exception of education, for which we use the latest vintage of the dataset originally introduced by Barro and Lee (2013). Details on variable definitions, transformations, and data sources are provided in Table A.1 and a list of IT countries used in this study in Table A.4 in the Appendix.

## 3.2 Identification of hawkishness

Our indicators of hawkishness and dovishness are based on a forward-looking specification of the Taylor rule with interest-rate smoothing. The central bank's target interest rate $r_t^*$ at time $t$ is given by:

$$r_t^* = b_{0,t}^* + b_{1,t}\big(E_t(\pi_{t,K}) - \pi_t^*\big) + b_{2,t}E_t(y_{t,K}), \tag{3}$$

where $\pi_t^*$ is the inflation target; $\pi_{t,K}$ and $y_{t,K}$ denote expected inflation and the expected output gap at horizon $K$; and $b_{0,t}^*$, $b_{1,t}$, and $b_{2,t}$ are time-varying policy parameters. The interest rate is gradually adjusted toward the target rate according to:

$$r_t = (1 - a_t)r_t^* + a_t r_{t-1} + \mu_t. \tag{4}$$

Rather than imposing an exogenously determined threshold, we classify a country as hawkish at time $t$ if its long-run response to inflation, $b_{1,t}$, exceeds the sample median.

Our approach follows Kim and Nelson (2006). Their two-step estimation strategy addresses two problems inherent to time-varying forward-looking models.

First, they address endogeneity by rewriting the Taylor rule in terms of future realizations – rather than expectations – and adding a bias-correction term.

When substituting Equation (3) into (4) and replacing expectations with future realizations, we obtain:

$$r_t = (1 - a_t)\big(b_{0,t} + b_{1,t}\pi_{t,K} + b_{2,t}y_{t,K}\big) + a_t r_{t-1} + e_t, \tag{5}$$

where

$$e_t = (1 - a_t)\Big[b_{1,t}\big(\pi_{t,K} - E_t(\pi_{t,K})\big) + b_{2,t}\big(y_{t,K} - E_t(y_{t,K})\big)\Big] + \mu_t. \tag{6}$$

The error term $e_t$ is correlated with the regressors $\pi_{t,K}$ and $y_{t,K}$ by construction, because both enter Expression (6). Recognizing that endogeneity arises from the forecast errors in inflation and the output gap, Kim and Nelson (2006) recommend augmenting Equation (5) with orthogonalized prediction errors for $\pi_{t,K}$ and $y_{t,K}$ generated by an auxiliary model.

This yields the following expression:

$$r_t = (1 - a_t)(b_{0,t} + b_{1,t}\pi_{t,K} + b_{2,t}y_{t,K}) + a_t r_{t-1} + \rho_1\sigma_{e,t}v^*_{1t} + \rho_2\sigma_{e,t}v^*_{2t} + \omega_t, \tag{7}$$

where $v^*_{1t}$ and $v^*_{2t}$ are the orthogonalized forecast errors and $\omega_t$ is the now independent error term.

Second, although the interest-rate-smoothing parameter generally falls between 0 and 1 in constant-parameter specifications, TVP models can misread increases in the policy rate as explosive dynamics, yielding estimates of $a_t$ above one. In this case, the long-run policy coefficients in Equation (5) are no longer well identified

By defining $a_t$ as

$$a_t = \frac{1}{1 + \exp(-b_{3,t})} \tag{8}$$

and modelling $b_{3,t}$ rather than $a_t$ itself as random-walk process addresses this issue.

However, the resulting model is no longer linear. To obtain a tractable representation, Kim and Nelson (2006) apply a first-order Taylor expansion to approximate the model linearly. The resulting linear specification, including the bias-correction terms described above, can then be estimated as a TVP model using a standard Kalman filter. In addition to the time-varying Taylor-rule parameters, their model also allows for heteroskedasticity modeled as stochastic volatility.

All Taylor-rules are estimated using quarterly data for the CPI and real GDP, where available, and unemployment otherwise. The data is sourced from the IMF and the Federal Reserve Economic Data FRED, hosted by the Fed St. Louis, and is seasonally adjusted using the X-13 ARIMA-SEATS. The output gap is computed using the Hodrick-Prescott filter.Where possible we use the policy rate, which is otherwise substituted by the lending rate. Where data availability allows, the samples used to estimate Taylor rules start in 1990.

# 4 Results

## 4.1 Time-Varying monetary policy reaction function

In almost all countries, the long-run response of monetary policy to inflation varies considerably over time (see Figure 1 - 7). Consistent with earlier evidence (see, e.g., El-Shagi and Ma; 2023), our estimates indicate that the long-run coefficients mostly do not satisfy the Taylor principle. Only in the Czech Republic and the United States do our point estimates exceed one throughout the sample. In 17 countries, the coefficient fluctuates around unity, while in the remaining 19 countries the Taylor principle is violated over the entire sample period.

The exact dynamics differ markedly across countries. While some, such as Costa Rica, India, Israel, Norway, Russia Federation, United Kingdom, and United States exhibit abrupt and sizable shifts, others display more gradual and continuous fluctuations.

The changes are slightly more pronounced during the period surrounding the global financial crisis (2008Q1–2012Q1). In roughly one quarter of the countries in our sample, we observe a considerable decline in the inflation response, suggesting a temporary shift toward greater dovishness. This pattern is consistent with the findings of Anderl and Caporale (2024) and Carvalho and Muinhos (2023), who likewise document that central banks tend to prioritize recovery over price stability in the aftermath of crises.

A caveat is that the confidence bounds are very wide for several countries. This reflects a common identification problem: the long-run inflation semi-elasticity of the interest rate is highly sensitive to the choice of smoothing parameter, with even minor adjustments producing substantial variation in the estimated long-run coefficient. To account for this issue, we implement a robustness test in our growth regressions by omitting all countries whose average standard error is above one.

## 4.2 Effects of key parameters of MPRF on economic growth

Across all specifications, the hawkishness dummy exerts a robust positive effect on growth. The effect is statistically significant under three of the four alternative estimators, see Table 1. The magnitude is economically meaningful, with estimated coefficients implying growth effects of 1.9 to 2.2 percentage points over a five-year horizon. This finding aligns with theoretical contributions (e.g., Gil et al.; 2023), which demonstrate that a strong policy reaction to inflation anchors expectations, lowers macroeconomic uncertainty, and promotes investment and long-run growth.

With the exception of education – which shows a negative association with growth – the control variables yield coefficients consistent with expectations. Such negative effects of human capital in country fixed-effects models, which absorb cross-country variation in education levels, are not unprecedented and are documented, for example, in Benhabib and Spiegel (1994) The coefficient on lagged log GDP is close to, yet below, unity, consistent with convergence dynamics documented in earlier work such as Barro et al. (2003).

Table 1: The effects of key parameters of the MPRF on economic growth

| Dependent var.: y | (1)<br>FE | (2)<br>sysGMM | (3)<br>diffGMM | (4)<br>Quasi-ML |
|---|---|---|---|---|
| $y_{t-1}$ | 0.7244*** | 0.9708*** | 0.7730*** | 0.6647*** |
| | (0.0651) | (0.0140) | (0.0407) | (0.1144) |
| $\text{Hawkish}_{t-1}$ | 0.0204* | 0.0223* | 0.0186*** | 0.0016 |
| | (0.0119) | (0.0121) | (0.0068) | (0.0225) |
| Gov. $\text{consumption}_{t-1}$ | -0.0154** | -0.0015 | -0.0083 | 0.0002 |
| | (0.0072) | (0.0020) | (0.0071) | (0.0091) |
| $\text{Investment}_{t-1}$ | -0.0029 | 0.0045*** | -0.0021 | -0.0114** |
| | (0.0021) | (0.0012) | (0.0020) | (0.0046) |
| Trade $\text{openness}_{t-1}$ | 0.0008 | 0.0003 | 0.0016* | 0.0003 |
| | (0.0009) | (0.0002) | (0.0008) | (0.0017) |
| $\text{Education}_{t-1}$ | -0.0247 | -0.0115 | -0.0310*** | -0.0038 |
| | (0.0181) | (0.0078) | (0.0090) | (0.0332) |
| Population $\text{growth}_t$ | -0.8211** | -0.6362** | -0.7036* | 0.2486 |
| | (0.4063) | (0.2452) | (0.4074) | (0.6304) |
| Constant | 3.1311*** | 0.3502*** | | 3.7034*** |
| | (0.6976) | (0.1049) | | (1.1620) |
| Observations | 94 | 94 | 57 | 44 |
| R-squared | 0.8815 | | | |
| Number of country_id | 37 | 37 | 32 | 19 |
| Country FE | YES | YES | YES | YES |
| Year FE | YES | YES | YES | YES |
| Arellano-Bond AR(1) | | 0.0698 | 0.0295 | |
| Arellano-Bond AR(2) | | 0.866 | 0.0961 | |
| Sargan | | 0.00280 | 0.0612 | |
| Hansen J | | 0.358 | 0.448 | |

*Note:* Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$
All variables measured in levels (including the hawkish) are end of period values. Lagged values therefore correspond to the value in the reference year of the current growth period. This avoids overlap between policy and its effects.

### 4.3 Robustness tests

**Remmoving high uncertainty countries** To ensure that our results are not driven by misclassified countries, we exclude all cases in which the average standard error of the long-run inflation coefficient exceeds unity. The effect of hawkishness on economic growth remains robust, with estimates slightly larger than in the baseline and ranging from 2.6 to 3.4 percentage points over a five-year horizon. This reflects that the positive baseline effect is primarily contributed by countries with precisely estimated inflation-response parameters (see Table 2).

**GDP in constant USD vs. GDP by PPP:** Consistent with the broader literature on the growth effects of central bank characteristics, such as inflation targeting and central bank independence, we measure growth using GDP in constant U.S. dollars. Yet the broader growth literature generally interprets economic growth as growth in PPP-adjusted real GDP. Replacing constant-USD GDP with PPP-adjusted GDP leaves the results broadly unchanged. The coefficients are somewhat larger than in the baseline, although they reach statistical significance only in the preferred difference-GMM estimator.

Jointly applying both robustness checks – substituting PPP-adjusted GDP and limiting the sample to observations with well-identified inflation semi-elasticities – raises the point estimates to 4.5 to 6.2 percent, and three of the four indicators remain statistically significant (see Table A.2 and A.3 in the appendix).

Table 2: The effects of key parameters of the MPRF on economic growth (Robustness test)

| Dependent var.: y | (1) FE | (2) sysGMM | (3) diffGMM | (4) Quasi-ML |
|---|---|---|---|---|
| $y_{t-1}$ | 0.8028*** | 0.9873*** | 0.8192*** | 0.5398*** |
| | (0.0687) | (0.0115) | (0.0644) | (0.1053) |
| $\text{Hawkish}_{t-1}$ | 0.0340** | 0.0264* | 0.0339*** | 0.0075 |
| | (0.0130) | (0.0151) | (0.0123) | (0.0223) |
| Gov. $\text{consumption}_{t-1}$ | -0.0134* | -0.0019 | -0.0115 | -0.0208* |
| | (0.0078) | (0.0038) | (0.0081) | (0.0116) |
| $\text{Investment}_{t-1}$ | -0.0101*** | 0.0089*** | -0.0108*** | -0.0245*** |
| | (0.0030) | (0.0016) | (0.0028) | (0.0061) |
| Trade $\text{openness}_{t-1}$ | 0.0017 | -0.0002 | 0.0022* | -0.0029 |
| | (0.0012) | (0.0003) | (0.0012) | (0.0017) |
| $\text{Education}_{t-1}$ | -0.0412** | -0.0270** | -0.0541*** | 0.0309 |
| | (0.0194) | (0.0106) | (0.0201) | (0.0316) |
| Population $\text{growth}_t$ | 0.0317 | -0.9120*** | 0.5585 | 0.6229 |
| | (0.5136) | (0.2795) | (0.8632) | (0.6892) |
| Constant | 2.5279*** | 0.1282 | | 5.6799*** |
| | (0.7199) | (0.1002) | | (1.1605) |
| Observations | 61 | 61 | 33 | 31 |
| R-squared | 0.9269 | | | |
| Number of country_id | 23 | 23 | 20 | 13 |
| Country FE | YES | YES | YES | YES |
| Year FE | YES | YES | YES | YES |
| Arellano-Bond AR(1) | | 0.0482 | 0.0304 | |
| Arellano-Bond AR(2) | | 0.877 | 0.978 | |
| Sargan | | 0.556 | 0.241 | |
| Hansen J | | 0.522 | 0.261 | |

*Note:* Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1
Countries with long-run inflation-response estimates whose standard errors exceed 1 in the TVP Taylor-rule are dropped. Such cases reflect weakly identified state-space parameters – usually from near-unit-root interest rate smoothing or short samples.

# 5 Conclusion

We analyze the relationship between the long-run inflation semi-elasticity of the policy rate and economic growth using data from 37 inflation-targeting countries. Our findings indicate sizable and statistically robust effects of a more hawkish policy stance – measured as an above-median inflation semi-elasticity – on economic growth, with annual impacts ranging from 0.5 percent to 1 percent across specifications.

Whereas the previous literature has largely focused on the growth effects of inflation targeting – typically interpreted as an official commitment to a stated inflation target – our results suggest that adopting the label alone may be insufficient. This underscores that even under monetary frameworks widely viewed as best practice, the time-inconsistency problem can still bias policymakers toward short-run stabilization at the cost of long-run growth.

This points to new directions for the literature on central bank institutions and growth: rather than examining individual institutional features and their interactions with external factors such as development or governance quality, it may be necessary to assess how different dimensions of monetary policy jointly interact. For example, the lack of emphasis on inflation – even within an inflation-targeting framework – may well be linked to insufficient central bank independence.

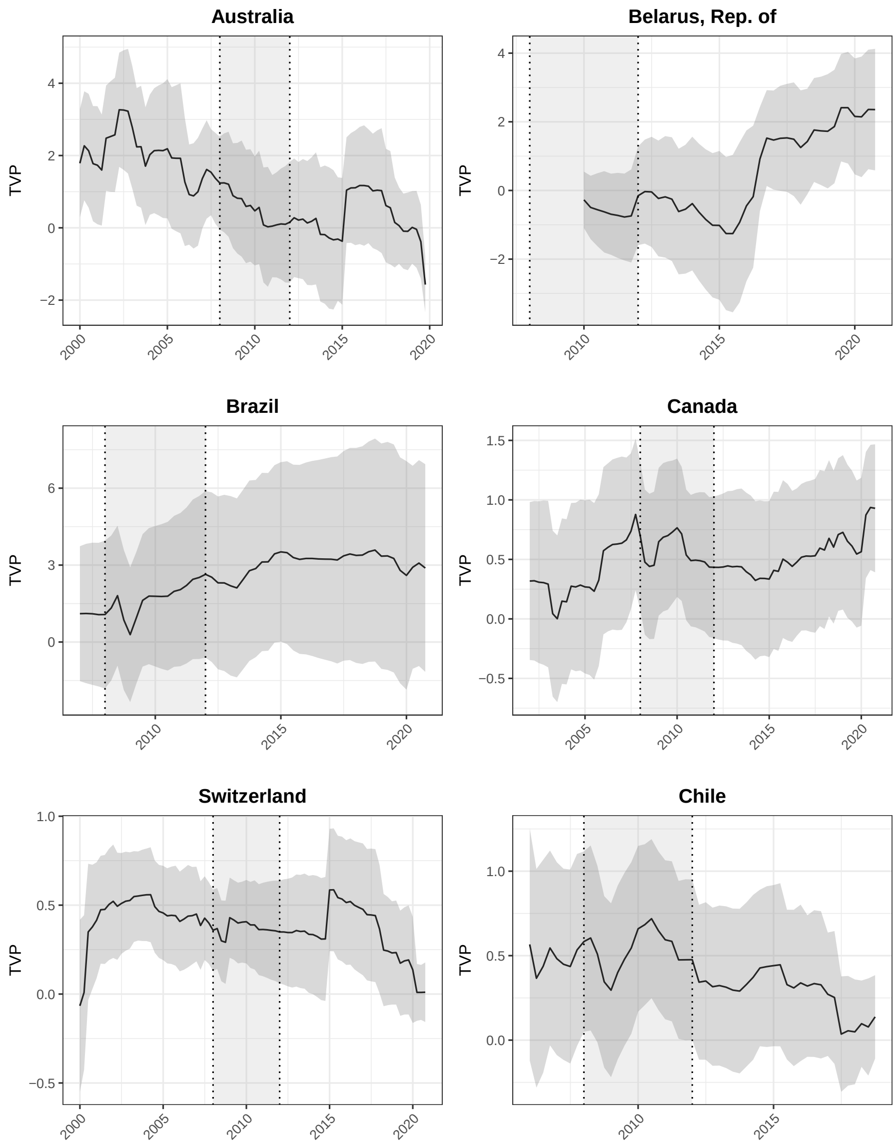


Figure 1: Time-varying response of policy rate to inflation and 90% confidence bands

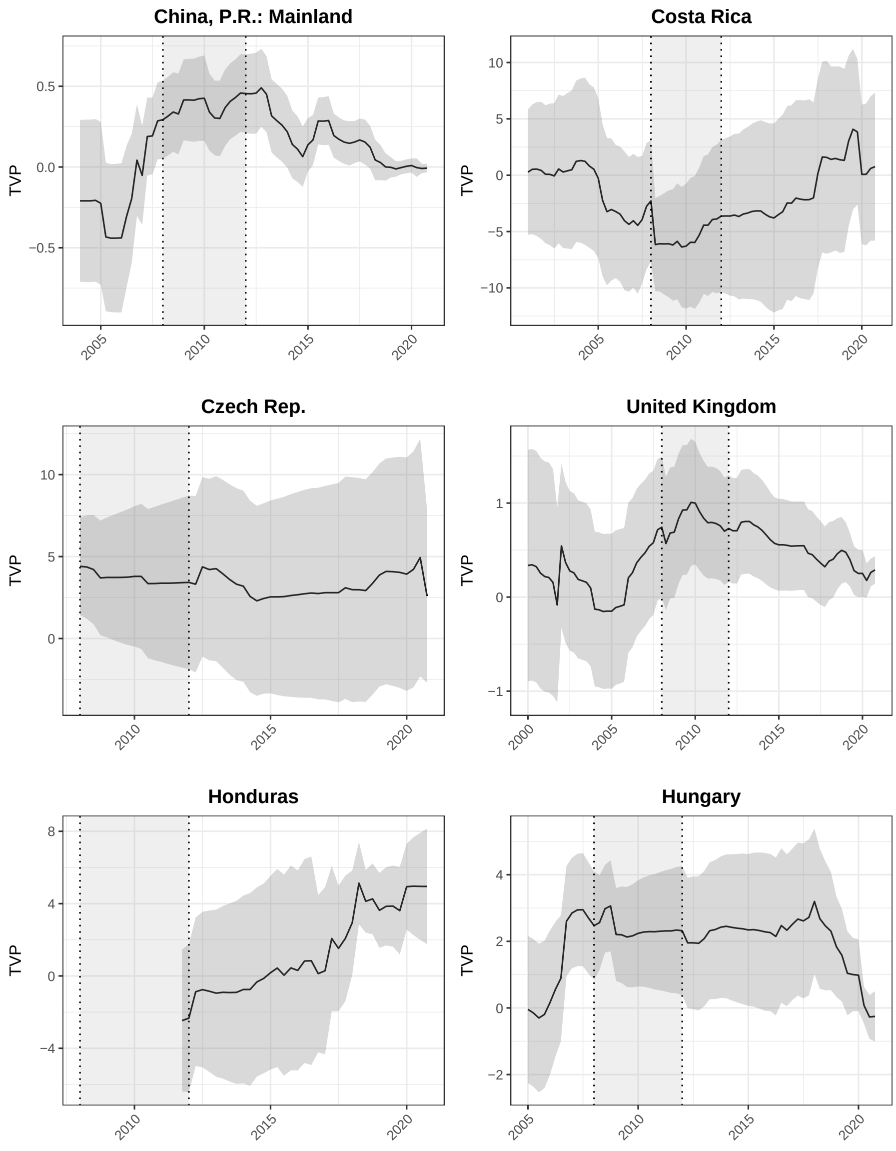


Figure 2: Time-varying response of policy rate to inflation and 90% confidence bands

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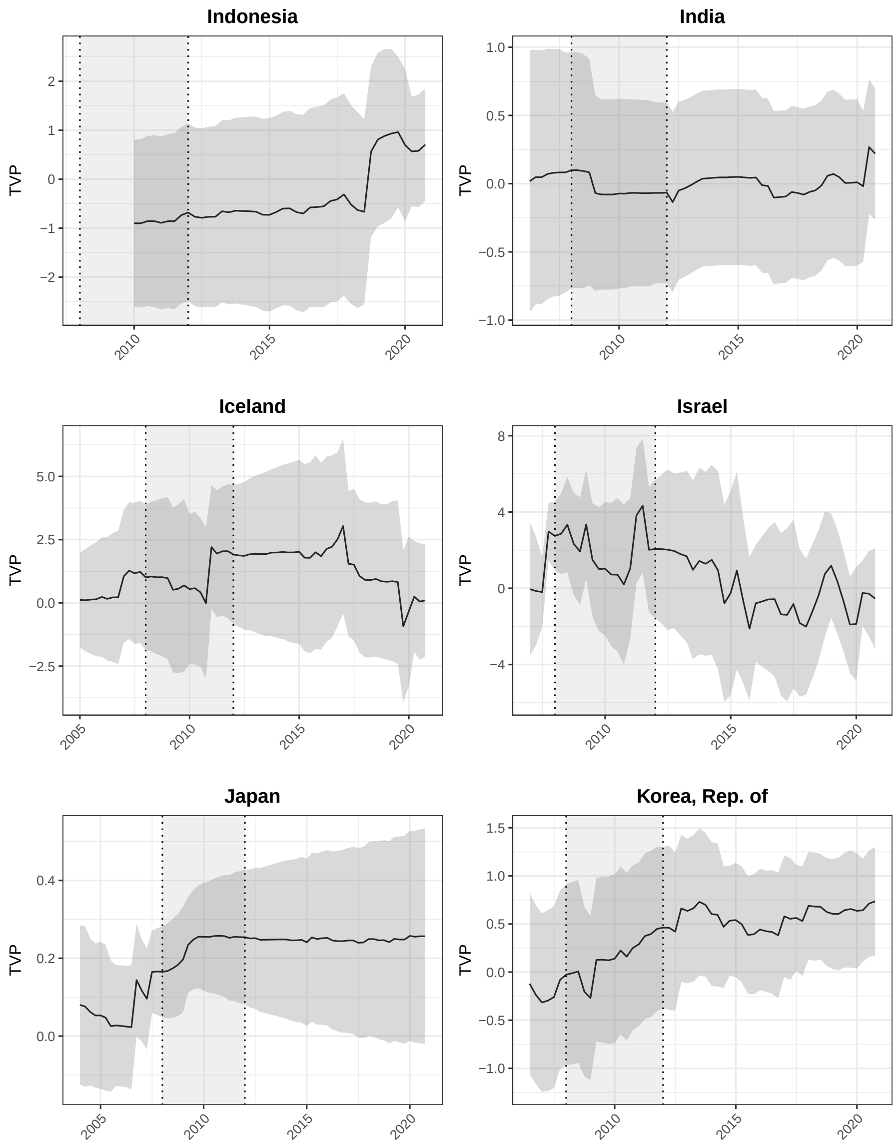


Figure 3: Time-varying response of policy rate to inflation and 90% confidence bands

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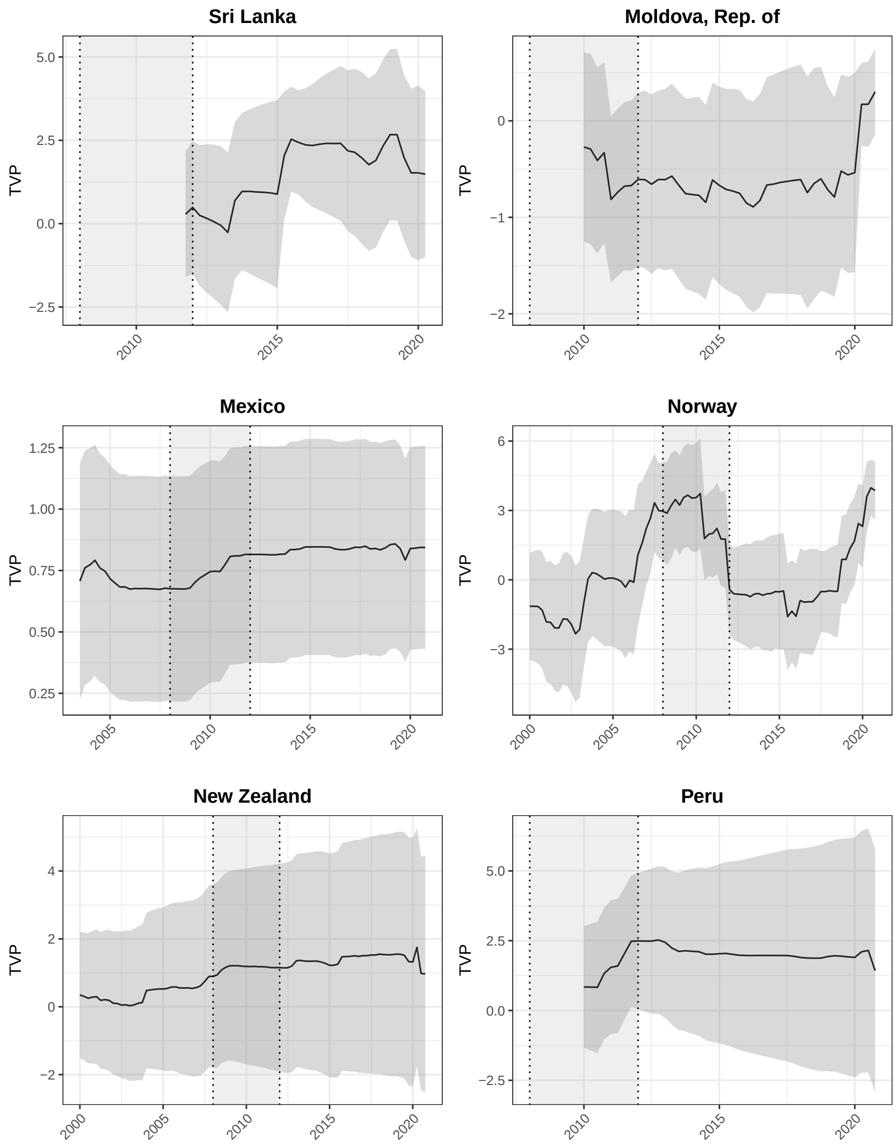


Figure 4: Time-varying response of policy rate to inflation and 90% confidence bands

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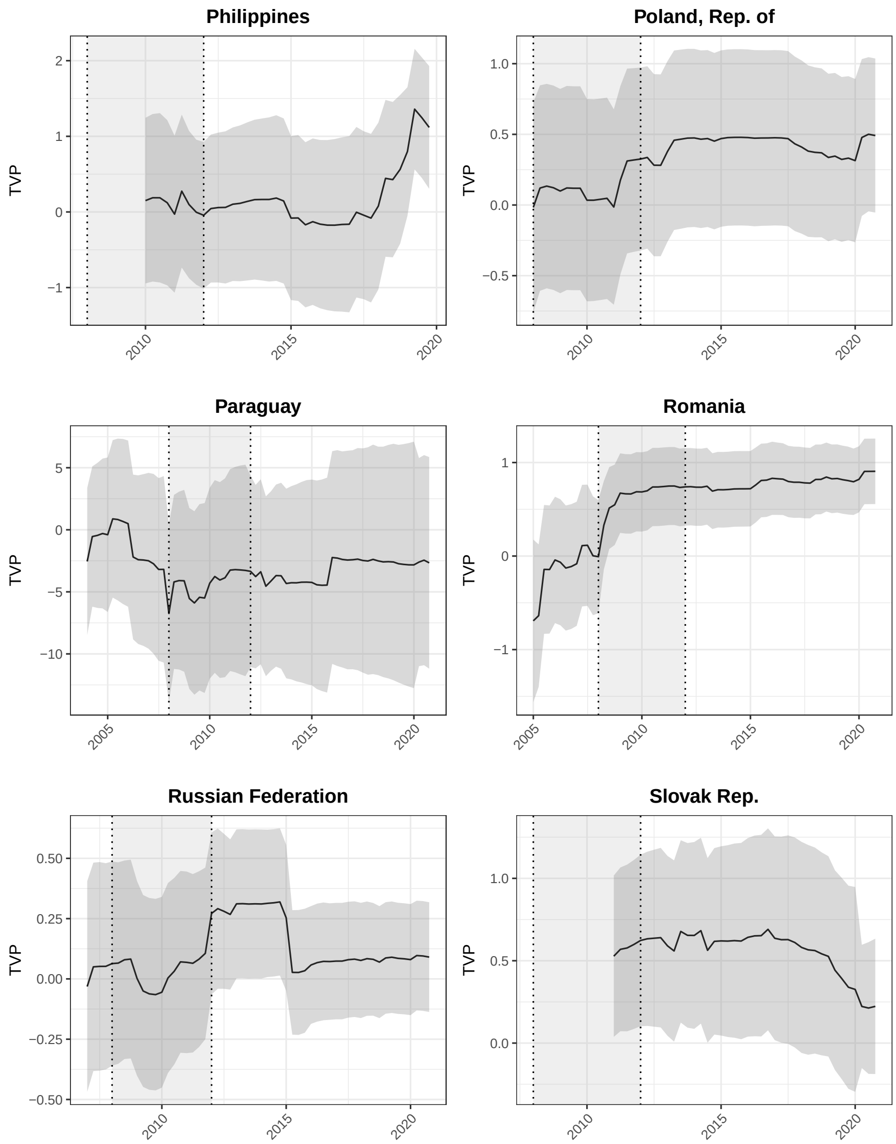


Figure 5: Time-varying response of policy rate to inflation and 90% confidence bands

.

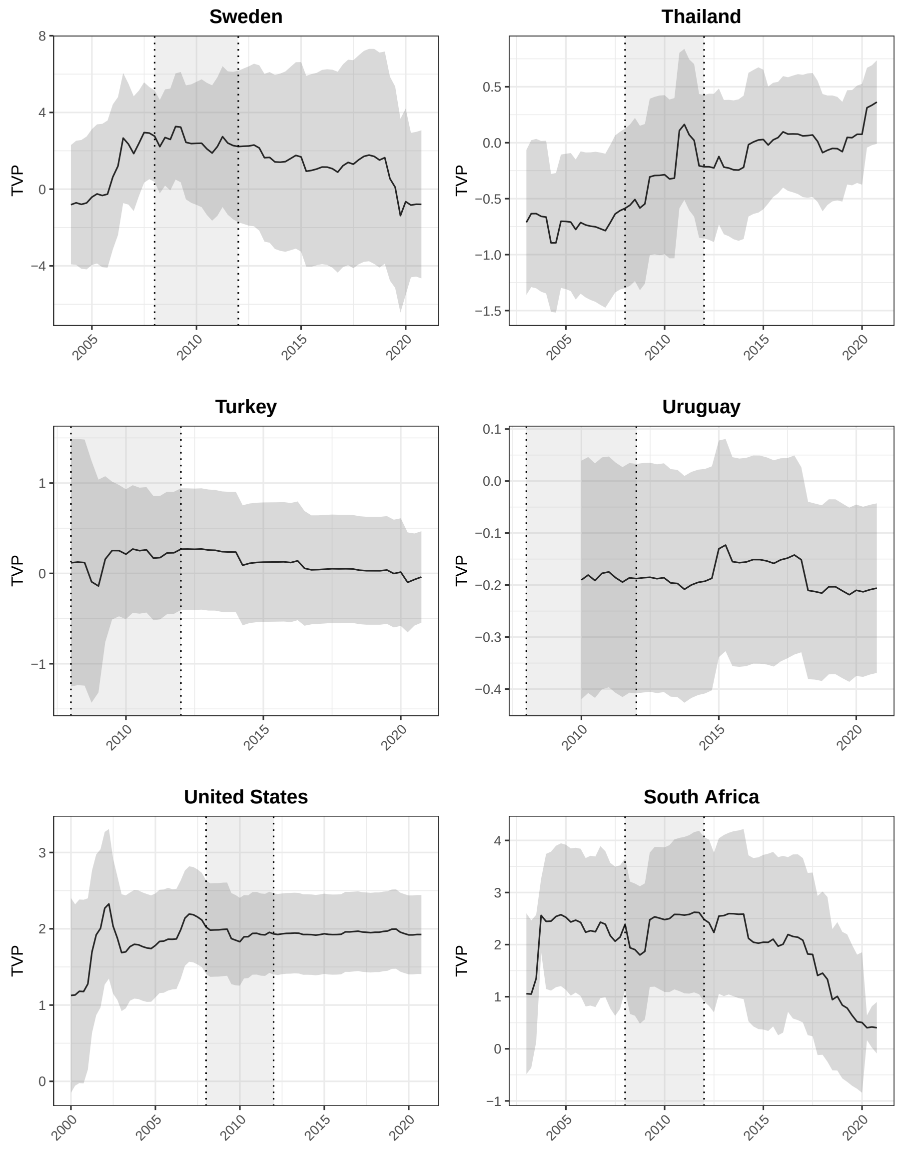


Figure 6: Time-varying response of policy rate to inflation and 90% confidence bands

.

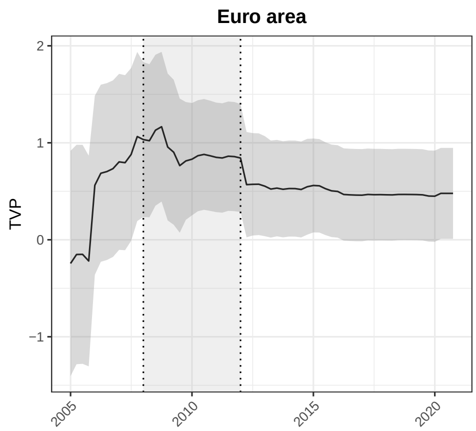


Figure 7: Time-varying response of policy rate to inflation and 90% confidence bands

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# 6 Appendix

Table A.1: Definition and sources of the variables used

| Area | Variable | Definition | Source |
|---|---|---|---|
| Monetary Policy: | | | |
| | Hawkish | Long-run inflation coefficient of MPRF | Be estimated from model (7) |
| | Interest rate | Lending rate, depositing rate or policy rate | IMF & FRED |
| | Inflation | Consumer Price Index (%) | IMF |
| | Real GDP | Gross Domestic Product, Real, Domestic Currency | IMF |
| Economic growth: | | | |
| | Real gdp per capita | GDP per capita (constant 2015 USD) | WDI |
| | Real GDP per capita | GDP per capita (adjusted for purchasing power parity (PPP)) | PWT 11.0 |
| | Gov. consumption | General government final consumption expenditure (% of GDP) | WDI |
| | Trade openness | Sum of imports and exports (% of GDP) | WDI |
| | Investment | Gross capital formation (formerly gross domestic investment) (% of GDP) | WDI |
| | Education | Secondary school attainment (% of population aged 15 and 64) | Barro and Lee (2013) |
| | Pop. growth | Change of total population in million | WDI |

Table A.2: The effects of key parameters of the MPRF on economic growth

| Dependent var.: y | (1) FE | (2) sysGMM | (3) diffGMM | (4) Quasi-ML |
|---|---|---|---|---|
| $y_{t-1}$ | 0.4432*** | 0.9988*** | 0.7143*** | 0.7383*** |
| | (0.0869) | (0.0192) | (0.1043) | (0.0842) |
| $\text{Hawkish}_{t-1}$ | 0.0287 | 0.0310 | 0.0495*** | 0.0346 |
| | (0.0193) | (0.0189) | (0.0142) | (0.0283) |
| Gov. $\text{consumption}_{t-1}$ | -0.0329*** | -0.0058** | -0.0265*** | -0.0130 |
| | (0.0108) | (0.0025) | (0.0079) | (0.0119) |
| $\text{Investment}_{t-1}$ | -0.0097*** | 0.0033 | -0.0122*** | -0.0181*** |
| | (0.0033) | (0.0025) | (0.0022) | (0.0058) |
| Trade $\text{openness}_{t-1}$ | -0.0032** | 0.0007 | -0.0021** | -0.0005 |
| | (0.0013) | (0.0006) | (0.0008) | (0.0022) |
| $\text{Education}_{t-1}$ | -0.0066 | -0.0240** | -0.0104 | 0.0451 |
| | (0.0273) | (0.0103) | (0.0289) | (0.0418) |
| Population $\text{growth}_t$ | 0.4211 | 0.3324 | 0.3486 | 1.7084** |
| | (0.6346) | (0.6301) | (0.4163) | (0.8082) |
| Constant | 8.5019*** | 0.2130 | | 4.0821*** |
| | (1.2355) | (0.3092) | | (1.0818) |
| Observations | 89 | 89 | 54 | 42 |
| R-squared | 0.9036 | | | |
| Number of country_id | 35 | 35 | 30 | 18 |
| Country FE | YES | YES | YES | YES |
| Year FE | YES | YES | YES | YES |
| Arellano-Bond AR(1) | | 0.00976 | 0.0194 | |
| Arellano-Bond AR(2) | | 0.393 | 0.696 | |
| Sargan | | 0.0118 | 0.339 | |
| Hansen J | | 0.0620 | 0.121 | |

*Note:* Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$
All variables measured in levels (including the hawkish) are end of period values.
Lagged values therefore correspond to the value in the reference year of the current growth period. This avoids overlap between policy and its effects.
Euro Area and Moldova are excluded from the PPP-adjusted GDP analysis because no PPP-adjusted GDP series exists for either of them.

Table A.3: The effects of key parameters of the MPRF on economic growth (Robustness test)

| Dependent var.: y | (1) FE | (2) sysGMM | (3) diffGMM | (4) Quasi-ML |
|---|---|---|---|---|
| $y_{t-1}$ | 0.4527*** | 0.9994*** | 0.7075*** | 0.6540*** |
| | (0.1102) | (0.0425) | (0.0901) | (0.0864) |
| Hawkish$_{t-1}$ | 0.0458* | 0.0287 | 0.0512*** | 0.0621** |
| | (0.0256) | (0.0278) | (0.0143) | (0.0289) |
| Gov. consumption$_{t-1}$ | -0.0410*** | -0.0043 | -0.0237* | -0.0415*** |
| | (0.0147) | (0.0040) | (0.0141) | (0.0149) |
| Investment$_{t-1}$ | -0.0123** | 0.0110*** | 0.0065 | -0.0365*** |
| | (0.0054) | (0.0034) | (0.0099) | (0.0080) |
| Trade openness$_{t-1}$ | -0.0034* | -0.0002 | -0.0024** | -0.0032 |
| | (0.0019) | (0.0011) | (0.0010) | (0.0024) |
| Education$_{t-1}$ | 0.0022 | -0.0178* | -0.0077 | 0.0778** |
| | (0.0344) | (0.0092) | (0.0264) | (0.0394) |
| Population growth$_t$ | 0.5718 | -0.6636 | 0.3558 | 2.3326*** |
| | (0.9733) | (0.9853) | (1.0104) | (0.8566) |
| Constant | 8.8472*** | 0.0082 | | 6.3782*** |
| | (1.6480) | (0.8059) | | (1.3926) |
| Observations | 58 | 58 | 36 | 29 |
| R-squared | 0.8988 | | | |
| Number of country_id | 22 | 22 | 19 | 12 |
| Country FE | YES | YES | YES | YES |
| Year FE | YES | YES | YES | YES |
| Arellano-Bond AR(1) | | 0.0494 | 0.0360 | |
| Arellano-Bond AR(2) | | 0.435 | 0.870 | |
| Sargan | | 0.476 | 0.173 | |
| Hansen J | | 0.439 | 0.401 | |

*Note:* Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$
Countries with long-run inflation-response estimates whose standard errors exceed 1 in the TVP Taylor-rule are dropped. Such cases reflect weakly identified state-space parameters – usually from near-unit-root interest rate smoothing or short samples.

Table A.4: List of inflation-targeting countries

| | | |
|---|---|---|
| Australia | Indonesia | Russian Federation |
| Belarus, Rep. of | Israel | Slovak Republic |
| Brazil | Japan | South Africa |
| Canada | Korea, Rep. of | Sri Lanka |
| Chile | Mexico | Sweden |
| China (P.R.: Mainland) | Moldova, Rep. of | Switzerland |
| Costa Rica | New Zealand | Thailand |
| Czech Republic | Norway | Turkey |
| Euro Area | Paraguay | United Kingdom |
| Honduras | Peru | United States |
| Hungary | Philippines | Uruguay |
| Iceland | Poland, Rep. of | |
| India | Romania | |