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RESEARCH ARTICLE / ARAŞTIRMA MAKALESİ

BANK AL-MAGHRIB AND THE MOROCCAN POLICY MIX : MONETARY REACTION ACROSS THREE CRISES (2008–2025)

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Abstract

This article examines Bank Al-Maghrib's monetary reaction function and the observed Moroccan policy mix across the 2008–2009 financial crisis, the 2020–2021 pandemic and the 2022–2023 inflationary shock. A smoothed Taylor rule uses quarterly observations from 2008Q1 to 2025Q2 with heteroskedasticity- and autocorrelation-consistent inference, alongside a descriptive annual policy-mix matrix. The benchmark indicates strong policy-rate inertia ($\rho \approx 0.95$) and positive associations with inflation and the output gap. Crisis dummies weaken the separate inflation association, while break diagnostics suggest instability around 2022. The implied long-run inflation response (approximately 1.23) is compatible with the Taylor principle; the bootstrap interval excludes zero but, like the delta-method interval, includes unity. Alternative output-gap measures preserve positive signs, with different magnitudes and sample sizes. Incomplete validation of the consumer price index (CPI) series before 2022 and the mechanical correction of a GDP discontinuity constrain inference. The annual policy configurations vary across crises and are more consistent with partly autonomous, shock-specific responses than with an identifiable coordination mechanism, although coordination through unobserved channels cannot be ruled out. The results motivate clearer policy communication, explicit assessment of fiscal–monetary trade-offs and better documentation of historical data.

Keywords : Smoothed Taylor Rule, Interest-Rate Smoothing, Monetary–Fiscal Policy Mix, Bank Al-Maghrib (Morocco), Emerging Economies.

JEL classification : E52, E58, E61, C32.

1. Introduction

Between 2008 and 2025, the Moroccan economy faces three distinct crises. The 2008–2009 financial crisis operates mainly through trade and external demand. The 2020–2021 pandemic combines an abrupt contraction in activity with a health emergency and exceptional fiscal support. The 2022–2023 inflationary shock shifts the policy priority towards price stability. In each

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episode, Bank Al-Maghrib (BAM) and the fiscal authorities respond to a common disturbance, but the co-movement of their instruments does not by itself establish institutional coordination.

It is precisely this articulation that is examined here. Describing how the central bank varies its policy rate in response to inflation and activity is not enough: one must also understand how that response combines with the fiscal stance, and whether the combination obeys a stable logic or is redefined with each shock. The literature offers solid footholds—the Taylor rule (Taylor, 1993), its extension to smoothing and forward-looking behaviour (Clarida et al., 1998, 2000; Rudebusch, 2002), and the policy-mix framework inherited from Tinbergen (1952) and Mundell (1962). Yet much of the available work tends to analyse the monetary rule, the fiscal stance, or the crisis episodes separately. This compartmentalisation leaves open a question that is simple to state but hard to settle: how do these dimensions fit together when shocks change in nature?

From this follows the article's research question: *does the Moroccan policy mix in the face of crises reflect an effective coordination between monetary and fiscal policy, or rather a sequence of partly autonomous responses, adjusted case by case to the nature of the shock?*

The analysis examines whether BAM's reaction function changes across crises and how the estimated monetary response relates to the observed fiscal stance. Strong policy-rate persistence is consistent with gradualism, but it can also reflect persistent omitted variables. The annual policy-mix configurations describe the combination of the two stances; they do not identify the institutional mechanisms behind that combination.

Morocco is, in this respect, an instructive emerging-market case. It combines a managed but gradually loosened exchange-rate regime, a central bank that builds credibility without operating a formal inflation target, a public-debt level that constrains fiscal space, and macroeconomic statistics whose quality improves unevenly over time. These are not idiosyncrasies of a single country but features shared, in varying proportions, by many small open economies that import a sizeable part of their inflation and remain exposed to external financial and commodity-price cycles. Studying how such an economy sequences monetary and fiscal action across three qualitatively distinct crises therefore speaks to questions of broad, cross-border relevance: how reaction functions behave when the dominant shock changes, how far a Taylor-type rule remains informative under weak transmission, and where the limits of mechanical rule-following lie.

The originality of the article lies less in estimating a Taylor rule—by now a classic exercise—than in using that rule as an instrument for interpreting the articulation of the Moroccan policy mix across three crises of differing nature. The contribution thus unfolds along four lines. Theoretically, the study brings together the smoothed Taylor rule and the Tinbergen–Mundell policy-mix framework for an emerging economy, where these grids are usually deployed in isolation. Empirically, it combines episode-based and endogenous break diagnostics with an annual policy-mix matrix covering three qualitatively different crises. Institutionally, it sheds light on the nature of the articulation of policies and draws implications for the conduct and communication of monetary policy. Methodologically, finally, it embraces a deliberate caution: the Moroccan data—only partial validation of the price index before 2022, a statistical break in gross domestic product (GDP) in 2014—require calibrating conclusions rather than forcing them.

The remainder is organised as follows. Section 2 situates the analysis within the literature. Section 3 presents the data, the stylised facts and the empirical approach. Section 4 presents the results, from the most robust to the most fragile. Section 5, the interpretive core of the article, draws the

economic lessons for the Moroccan policy mix and discusses the limitations as conditions of interpretation. Section 6 concludes.

2. Theoretical Framework and Literature Review

2.1. Taylor rule, interest-rate smoothing and the Taylor principle

The idea that a central bank's behaviour can be summarised by a simple rule goes back to Taylor (1993): the nominal rate responds to the deviation of inflation from target and to the output gap. The Taylor principle extends this intuition with a stability condition—a more-than-proportional response to inflation, that is, a long-run inflation coefficient greater than one. This formulation appears too bare. Clarida et al. (1998, 2000) add two standard ingredients: rate smoothing, which captures central banks' reluctance to make abrupt moves, and a forward-looking dimension, oriented toward expected inflation. Rudebusch (2002), however, cautions that strong estimated inertia is not necessarily a sign of deliberate smoothing: it may also reflect the persistence of omitted variables. This warning is valuable for the Moroccan case, where the autoregressive coefficient is high and calls for a cautious reading.

Underneath the appeal of rules lies a deeper rationale that bears directly on emerging-market central banking. The time-inconsistency literature (Barro & Gordon, 1983; Kydland & Prescott, 1977) shows that a discretionary authority, tempted to exploit short-run output gains, ends up delivering higher average inflation once private agents anticipate that temptation; commitment to a rule-like, predictable conduct, and the reputation that sustains it, can improve outcomes. Read in this light, the smoothing and the more-than-proportional response that the Taylor principle requires are not merely technical features of an estimated equation: they are devices through which a central bank signals consistency and anchors expectations. For an institution that builds credibility gradually and without a formal numerical target, this reputational dimension matters at least as much as the precise size of any single rate move.

Two cautions follow for the empirical exercise. First, the long-run inflation coefficient, not the contemporaneous one, is what speaks to the Taylor principle, and it is estimated with considerable imprecision once strong inertia inflates the standard errors of the implied ratio. Second, applying the Taylor principle as a pass/fail verdict in an emerging economy is hazardous: imperfect transmission, an exchange-rate objective and noisy real-time data all blur the mapping from the estimated rule to the underlying behaviour. The principle is therefore best treated as an interpretive benchmark rather than as a norm to be mechanically enforced.

2.2. Policy mix, fiscal–monetary coordination and crises

The policy mix is the art of combining the two main levers of economic policy. Tinbergen (1952) sets out its grammar: pursuing several independent objectives requires at least as many instruments, and hence a deliberate assignment of the levers. Mundell (1962) sharpens the logic by assigning each authority the objective on which it holds a comparative advantage. The subsequent literature formalises the interaction between authorities in terms of "active" and "passive" regimes (Leeper, 1991), showing that macroeconomic stability depends on the combination of monetary and fiscal behaviour more than on either taken in isolation. From this work follows a useful typology: depending on whether the two stances reinforce, ignore or counteract one another, the policy mix may be described as coordinated, neutral or conflictual.

Two implications of this framework guide the interpretation below. First, a regime in which the fiscal stance is "passive"—adjusting so as not to obstruct the monetary anchor—supports price stability, whereas a configuration in which fiscal pressure dominates can force monetary policy to accommodate, with inflationary consequences. Second, and crucially for what follows, the observation that the two stances move in the same direction is not, by itself, evidence of an organised coordination mechanism. Crisis periods put this distinction to the test, since the superposition of real, financial and price shocks can produce a concordance of stances without any institutionalised coordination existing for all that. Keeping apart genuine coordination from mere circumstantial concordance is one of the interpretive tasks the empirical analysis must shoulder.

2.3. Emerging economies and the Moroccan case

Transposed to an emerging economy, these frameworks meet well-identified frictions. Monetary transmission is often incomplete, the exchange-rate regime weighs on the room for manoeuvre, public debt limits fiscal space, and cyclical statistics are of uneven quality. The reaction functions estimated for these economies indeed display a high weight of smoothing and a marked sensitivity to the exchange rate (Mohanty & Klau, 2004), consistent with the idea that the exchange rate is both an implicit objective and a source of shocks (Taylor, 2001). A recurrent finding of this literature is that part of the inflation observed in small open economies is imported rather than demand-driven, so that a domestic policy rate can do little, in the short run, against a terms-of-trade or energy-price shock—which makes the credibility of the announced path, rather than the amplitude of any single move, the decisive margin.

Morocco offers a nuanced illustration: long-contained inflation, a basket peg gradually loosened, significant public debt, and constant attention to financial stability. In such an environment, the credibility of the announced path sometimes matters more than the size of any isolated change in the rate. This is why the Taylor rule is used here as a reading grid, not as a norm against which to assess the central bank mechanically. On the technical side, the analysis draws on the standard time-series toolkit—stationarity testing, output-gap filtering (Hamilton, 2018; Hodrick & Prescott, 1997), bounds cointegration (Pesaran et al., 2001), structural-break tests (Bai & Perron, 2003) and robust inference (Newey & West, 1987)—and on the institutional and statistical framework provided by Bank Al-Maghrib, the Haut-Commissariat au Plan (HCP) and the International Monetary Fund (IMF).

Three features of this setting shape what an estimated rule can and cannot reveal, and they recur across emerging economies. First, when the exchange rate is managed and partly anchors prices, the policy rate carries part of an external-balance objective that a domestic reaction function does not display explicitly, so that an apparently muted response to inflation may conceal an active defence of the nominal anchor. Second, where public debt is sizeable, the interest rate is also a fiscal variable, and the central bank must weigh the debt-service consequences of tightening against its disinflationary benefits—a tension that pushes towards gradualism quite apart from any preference for smoothing. Third, when cyclical statistics are revised substantially, as the 2014 national-accounts rebasing illustrates, the real-time output gap on which decisions actually occur differs from the one reconstructed ex post, so that estimated slope coefficients mix genuine behaviour with measurement. These three considerations do not weaken the case for estimating a reaction function; they define the interpretive caution with which its coefficients must be read, and they motivate the battery of robustness checks reported below.

Recent evidence sharpens the comparison. For Morocco, Ouchchikh et al. (2025) use wavelet methods and find uneven monetary transmission, with evidence in favour of the credit channel and more limited investment responses through interest rates. Hammou Ou Ali (2025), using a structural vector autoregression (SVAR) over 2006–2024, reports a delayed decline in housing prices following monetary tightening. These studies examine transmission rather than estimate the same reaction equation; they motivate attention to financial conditions and adjustment lags without validating the coefficients reported here.

For the BRICS, Waheed et al. (2024) estimate a Bayesian dynamic stochastic general equilibrium (DSGE) model and document country-specific effects of credit spreads and fiscal imbalances on macroeconomic outcomes. For South Africa, Liu and Mustapher (2025) examine US monetary spillovers in a small-open-economy model with a fiscal block and show that outcomes depend on the monetary–fiscal regime. Their structural analysis distinguishes mechanisms that the present descriptive policy-mix matrix cannot identify. Together, these contributions motivate a crisis-sensitive reading of the Moroccan reaction function while limiting comparisons to the questions and assumptions each design can address.

3. Materials and Methods

3.1. Data and stylised facts

Four quantities suffice to tell most of this story, and each plays a distinct role. The policy rate is the central bank's principal instrument of signalling and action. Inflation is the price-stability signal—that is, the primary objective to which this signal is meant to respond[3]. Real GDP, and the output gap derived from it[4], indicate the economy's position in the cycle. The fiscal stance, finally, embodies the second pillar of the policy mix, whose articulation with the first is the object of the study. Table 1 sets out the definitions and sources.

Table 1. Description of variables and sources.

Variable	Definition	Frequency	Source
Policy rate	Bank Al-Maghrib key policy rate (%)	Quart. / Ann.	BAM
Inflation	Year-on-year log change in the CPI (base 2017 = 100)	Quart.	HCP
Real GDP	GDP at constant prices (MAD millions)	Quart.	HCP
Output gap	Filtered GDP (HP $\lambda = 1600$), log corrected for the 2014 break	Quart.	Authors' calc.
Fiscal stance	Fiscal impulse (pp); budget balance (% of GDP)	Ann.	IMF / Authors' calc.

Source. *Compiled by the authors from BAM (n.d.), HCP (n.d.), and IMF (2023, 2024). CPI: consumer price index; pp: percentage points; MAD: Moroccan dirham.*

[3]The inflation measure used is observed inflation, computed as the year-on-year log change in the CPI, rather than a deviation from an explicit numerical target. Bank Al-Maghrib does not pursue a formal inflation target over most of the period studied.

[4]The Hodrick–Prescott filter is set to $\lambda = 1600$, the usual value for quarterly data. The output gap is expressed as 100 times the difference between corrected log real GDP and its filtered trend.

The analysis rests on two datasets built from a single source file. The series are drawn from Bank Al-Maghrib (n.d.) for the policy rate, the Haut-Commissariat au Plan (n.d.) for inflation and GDP, and the International Monetary Fund (2023, 2024) for the budget balances. The BAM and HCP series come from the institutions' online databases (source access verified: 17 September 2026) and include observations through 2025Q2; the IMF sources provide the annual budget balances through 2023. The quarterly dataset spans 2007Q1–2025Q2, with the main estimation running over 2008Q1–2025Q2 (70 observations); the annual dataset, which feeds the policy-mix matrix, covers 2008–2023. Inflation is measured as the year-on-year log change in the price index.

Two data limitations condition the interpretation. First, the assembled GDP series displays a discontinuity of approximately +15.6% between 2013Q4 and 2014Q1, which corresponds to the change in the base year of the national accounts. The adjustment described below removes this level discontinuity before filtering; it does not reconstruct an official backcast series. Second, internal validation of the CPI observations in the working dataset is complete only from 2022 onward. This is a limitation of the dataset used here, not evidence that the official HCP CPI is generally unreliable.

The output-gap construction starts from $x_t = \ln(Y_t)$, where Y_t is real GDP. As shown in Figure S3, the corrected series applies $x^*_t = x_t - \ln(1.156) \times 1\{t \geq 2014Q1\}$. Thus, the post-break level is divided by 1.156; within-segment growth rates remain unchanged. This mechanical splice treats the observed jump as a level discontinuity. It may also remove genuine growth at the junction, and it does not correct changes in coverage, sector weights or seasonal treatment. The estimated break near 2014 therefore requires particular caution.

The benchmark applies the two-sided Hodrick–Prescott (HP) filter with $\lambda = 1{,}600$ to the corrected log-GDP series. The trend τ_t minimises $\Sigma_t(x^*_t - \tau_t)^2 + 1{,}600\ \Sigma_t[(\tau_{t+1} - \tau_t) - (\tau_t - \tau_{t-1})]^2$, and the output gap is $100(x^*_t - \tau_t)$, an approximate percentage deviation from the statistical trend (Hodrick & Prescott, 1997). Figures S2–S3 show the resulting gap and level adjustment. The two-sided filter uses future observations and is sensitive to endpoints, especially around the pandemic and the end of the sample. It therefore does not measure a real-time policy information set.

One-sided HP and Hamilton-based alternatives assess sensitivity to trend extraction (Hamilton, 2018; Tables S12–S13). Their reported samples contain 59 and 63 observations, respectively, compared with 70 in the benchmark. The available outputs do not establish a common-sample comparison, so stable coefficient signs are evidence of limited sensitivity rather than proof of filter invariance.

The three crises are identified by dummy variables (Table 2), which then make it possible to test whether the reaction function shifts from one regime to another.

Table 2. Crisis coding.

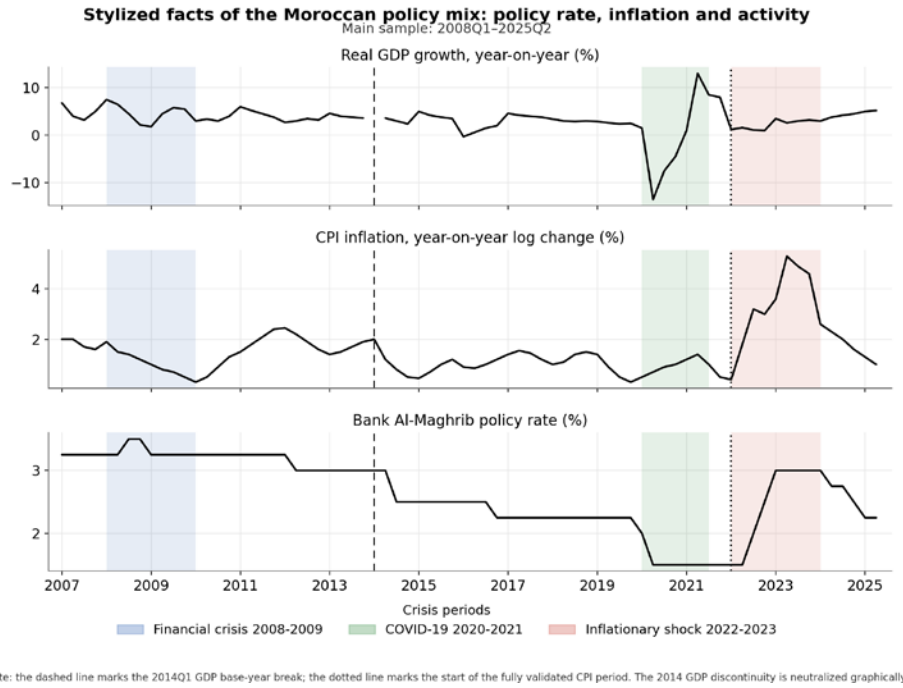
Crisis	Period (quarters)	Nature
Financial crisis	2008Q1–2009Q4	Financial and external-demand shock
COVID-19	2020Q1–2021Q2	Supply and demand shock, lockdown
Inflationary shock	2022Q1–2023Q4	Price surge and monetary tightening

Source. *Authors' coding.*

Figure 1 captures the point. The policy rate traces a staircase line that moves only rarely and in small steps, whereas activity undergoes considerable swings—the 2020 collapse and rebound bear

witness—and the 2022–2023 inflation reaches levels without equivalent over the period. This contrast between a heavily smoothed instrument and a highly volatile environment is the first stylised fact the analysis must explain.

Figure 1. Policy rate, inflation and real GDP growth, 2007Q1–2025Q2.



Note. *Shaded bands mark the three crisis episodes; the dashed line marks the 2014Q1 GDP base-year break; the dotted line marks 2022Q1, the start of the fully validated CPI period in the working dataset. Source: authors' calculations from BAM (n.d.) and HCP (n.d.).*

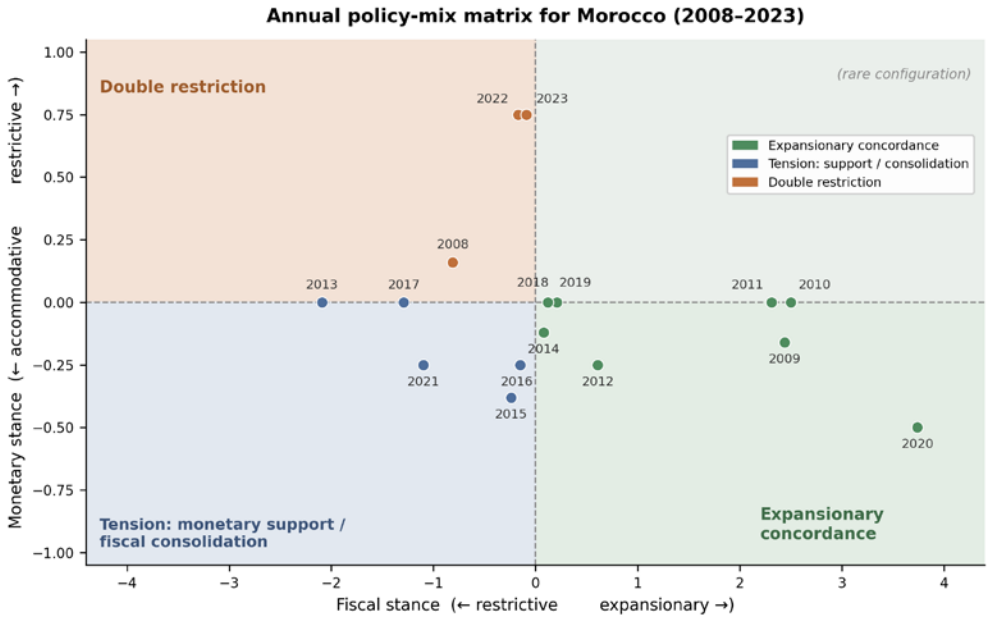
A second stylised fact concerns the rhythm of the rate itself. Over the whole sample the policy rate moves on only a handful of occasions and almost always in increments of a quarter of a percentage point, with longer plateaux between adjustments than is typical of central banks operating a formal inflation target. The downward drift of the late 2000s and early 2010s, the sharp easing of 2020 and the tightening of 2022–2023 stand out precisely because they interrupt these plateaux. Read together, the two stylised facts—a smoothed instrument and a turbulent environment—already foreshadow the central tension of the analysis: an institution that prizes predictability is repeatedly confronted with shocks that would, on a mechanical reading of any rule, call for larger and faster moves than those actually observed.

To make the articulation of the two policies visible, we set against one another, year by year, the monetary stance (change in the policy rate) and the fiscal stance (fiscal impulse)[5]. Figure 2 organises this comparison and brings out three recurring configurations: phases of expansionary concordance, where both levers support the economy; episodes of tension, where monetary support coexists with fiscal consolidation; and rare moments of double restriction. It bears

[5]The fiscal impulse is defined as the annual change in the budget balance, in percentage points of GDP, computed from IMF data. A positive value—an improvement in the balance—signals a restrictive stance; a negative value—a deterioration—signals an expansionary stance. In the matrix, the sign is reversed so that fiscal expansion points to the right.

emphasising that this matrix is a descriptive reading device[6]. On such a short annual sample, it helps to characterise the observed configurations without ever establishing causality between the two policies.

Figure 2. Annual policy-mix matrix for Morocco (2008–2023).



Note. *Position of each year according to the change in the policy rate and the fiscal impulse (sign reversed: fiscal expansion points to the right). Detailed annual classification is provided in the supplementary material. Source: authors' calculations from BAM (n.d.) and IMF (2023, 2024).*

3.2. Empirical approach

The aim is not to impose a mechanical rule on Moroccan monetary policy, but to use the smoothed Taylor rule as a lens through which to read Bank Al-Maghrib's behaviour in the face of shocks of differing nature. This nuance governs the entire approach: the estimated coefficients are treated as convenient summaries of behaviour, not as proof that a norm is being followed or breached.

A clarification of scope is in order at the outset. The econometric analysis primarily identifies Bank Al-Maghrib's monetary reaction function. The fiscal dimension, for its part, is used in a descriptive annual reading of the policy mix: it serves to characterise the observed configurations without establishing a causal relationship between the two authorities. This is why the vocabulary used carefully distinguishes the estimated reaction function, on the one hand, from the observable configuration of the policy mix, on the other. Box 1 makes this division of labour explicit, so that the reach of each claim can be judged against what the data can actually bear.

Box 1. Scope and limits of inference
The analysis describes a persistent monetary reaction function, specification-dependent inflation and activity associations, candidate breaks and annual policy-mix configurations. It does not establish causal monetary–fiscal coordination, optimal gradualism, a definitive Taylor-principle verdict or structurally identified VAR effects.

[6] The matrix is strictly descriptive: it characterises observable annual configurations and establishes no causal relationship between the fiscal and monetary stances.

The benchmark specification relates the policy rate to its own lagged value—the smoothing term—to inflation and to the output gap. In the absence of expectations series available over the whole period, it is kept in contemporaneous form, with inflation and the output gap dated at t. In its theoretical form, it is written

$$i_t = \rho i_{t-1} + (1 - \rho)(c + \beta\pi_t + \gamma gap_t) + \delta' D_t + \varepsilon_t, \quad (1)$$

and, in the linearised form actually estimated,

$$i_t = \alpha + \rho i_{t-1} + \varphi\pi\pi_t + \varphi gap\, gap_t + \delta' D_t + \varepsilon_t. \quad (2)$$

Here $\alpha = (1 - \rho)c$, $\varphi\pi = (1 - \rho)\beta$ and $\varphi gap = (1 - \rho)\gamma$; D_t denotes the vector of crisis dummies (Table 2) and is omitted in the benchmark model. The long-run inflation response is $\beta LT = \varphi\pi/(1 - \rho)$, conditional on a stable specification and $|\rho| < 1$. Ordinary least squares (OLS) estimates use Newey–West heteroskedasticity- and autocorrelation-consistent (HAC) standard errors with four lags (Newey & West, 1987). The delta method and 2,000 residual bootstrap replications summarise uncertainty around βLT.

Crisis dummies allow intercept shifts at the dates specified in Table 2; interactions with inflation and the output gap allow the corresponding slopes to differ. These episode-based specifications therefore use predetermined crisis windows. Separately, the Bai–Perron procedure (Bai & Perron, 2003) and the sup-F test (Andrews, 1993) assess unknown structural breaks. The OLS-based cumulative sum (CUSUM) diagnostic (Ploberger & Krämer, 1992) assesses cumulative parameter instability; it does not estimate break dates. The tests address different alternatives, so their conclusions need not coincide.

Robustness checks distinguish evidence on a long-run levels relationship from short-run adjustment within the autoregressive distributed lag (ARDL) error-correction model (ECM). Additional checks use alternative output-gap measures, a simple AR(1) benchmark and a restricted CPI subsample (2021Q4–2025Q2). A vector autoregression (VAR) with the recursive ordering inflation → output gap → policy rate provides a conditional description of the dynamics. The responses depend on this Cholesky ordering. Tables S10–S18 provide the detailed diagnostics.

4. Results

4.1. A highly gradual central bank

The benchmark smoothing coefficient is approximately 0.95 (Table 3). The current policy rate is therefore closely associated with its lagged value, conditional on inflation and the output gap. This persistence is consistent with gradual adjustment, but the regression does not distinguish deliberate smoothing from persistent omitted determinants or infrequent changes in the policy instrument.

This gradualism has a statistical consequence that must be named at once, lest one misread what follows: the models' very high goodness of fit stems first from it. Section 4.5 returns to the point, but it bears stating now—explaining today's rate by yesterday's is, mechanically, highly effective.

4.2. Inflation matters, but its weight depends on the regime

In the simple specification, the result accords with economic intuition: the central bank responds in the right direction to both inflation ($\varphi_\pi \approx 0.059$; p = 0.034) and the output gap ($\varphi_{gap} \approx 0.024$; p < 0.001). The rate rises when prices accelerate and when activity tightens—behaviour expected of an authority concerned with stability (Table 3, column 1).

With crisis dummies, inertia remains high and the output-gap coefficient remains significant, whereas the inflation coefficient is no longer statistically significant (Table 3, column 2). Section 4.3 examines the overlap between inflation and the inflationary-shock dummy. This overlap weakens separate identification of their contributions; it does not demonstrate either the presence or the absence of a structural inflation response. The improvement in fit remains modest, and the Bayesian information criterion (BIC) favours the parsimonious specification.

Table 3. Estimation of the smoothed Taylor rule (OLS, Newey–West HAC errors).

Variable	Simple Taylor	Taylor with crises
i_{t-1} (ρ)	0.952*** (0.030)	0.924*** (0.046)
π_t (φ_π)	0.059** (0.027)	0.020 (0.031)
gap_t (φ_{gap})	0.024*** (0.006)	0.021*** (0.007)
Constant	0.012 (0.066)	0.125 (0.103)
Financial crisis	—	0.094* (0.055)
COVID-19 crisis	—	−0.058 (0.078)
Inflationary shock	—	0.169* (0.086)
N	**70**	**70**
Adjusted R^2	**0.951**	**0.954**
AIC / BIC	**−82.75 / −71.51**	**−85.34 / −67.35**

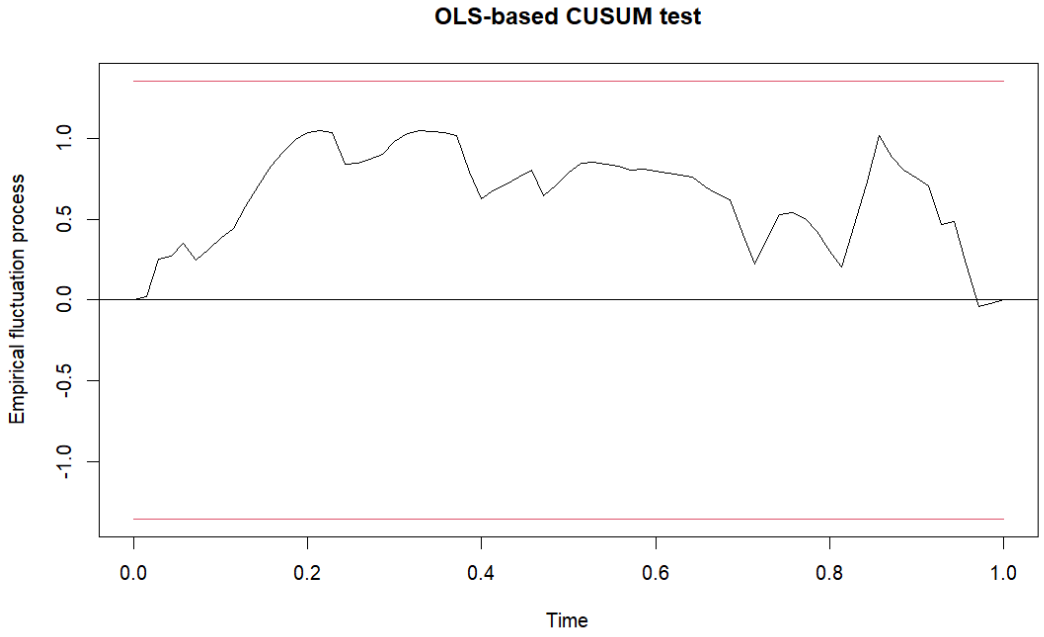
Note. *HAC standard errors in parentheses. *** p < 0.01; ** p < 0.05; * p < 0.10. AIC: Akaike information criterion; BIC: Bayesian information criterion. Detailed coefficients, variance inflation factors and crisis interactions are reported in the supplementary material.*

4.3. Evidence of instability around the inflationary episode

Inflation and the 2022–2023 dummy overlap substantially (correlation of 0.79; Table S8). In the model containing inflation and only the inflationary-shock dummy, the inflation p-value is 0.495; with all three crisis dummies it is 0.510. Both exceed conventional significance thresholds. The evidence indicates limited ability to distinguish an inflation association from an episode effect within this small sample.

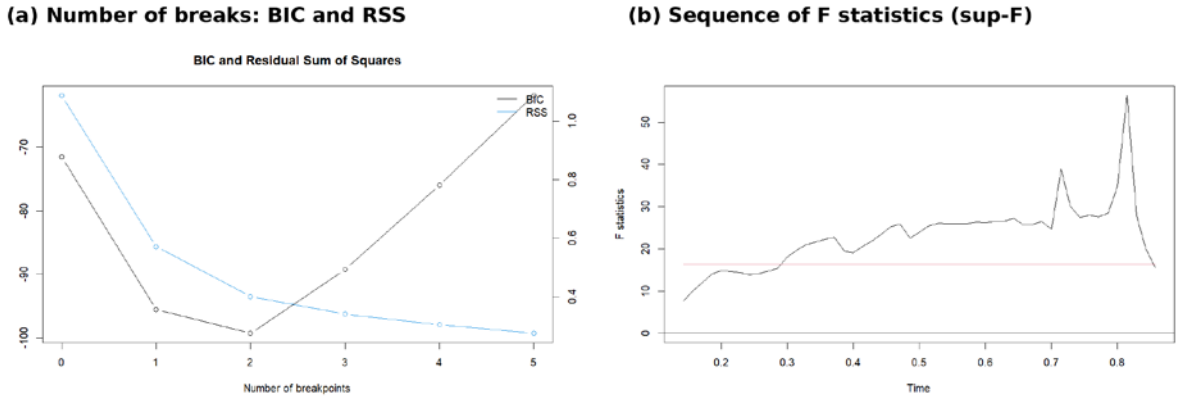
Bai–Perron estimates place candidate breaks around 2014Q2 and 2022Q1 (Table S9; Figure 4a). The first lies close to the 2014 national-accounts rebasing; the present data do not allow us to determine whether it reflects measurement or a change in policy behaviour. The second coincides with the start of the coded inflationary episode; it should not be equated with the date of the first policy-rate increase. The sequence of F statistics exceeds the 5% sup-F critical value, with a maximum close to 2022Q1 (Figure 4b), whereas the OLS-based CUSUM path remains within its 5% boundaries (Figure 3). Non-rejection by CUSUM is not proof of constant parameters and does not invalidate tests directed at specific breaks.

Figure 3. OLS-based CUSUM test of parameter stability.



Note. *OLS-based CUSUM process of the benchmark smoothed Taylor rule (Table 3, column 1). The horizontal lines at ±1.36 are the 5% critical boundaries (Ploberger & Krämer, 1992); the horizontal axis is normalised sample time. The path stays within the boundaries, so the test does not reject parameter constancy. Source: authors' calculations (also Figure S7).*

Figure 4. Structural-break diagnostics.



Note. *Panel (a) plots the Bayesian information criterion (BIC) and the residual sum of squares (RSS) against the number of breakpoints; the BIC reaches its minimum at two breaks (2014Q2 and 2022Q1; Table S9). Panel (b) plots the sequence of F statistics over normalised sample time (15% trimming) with the 5% critical boundary of the sup-F test (Andrews, 1993); the peak lies close to 2022Q1. Source: authors' calculations (also Figures S5–S6).*

Taken together, the results suggest a change in the estimated reaction function around the inflationary episode. They do not isolate a structural change in BAM's preferences from the effects of imported inflation, omitted variables and data limitations.

4.4. The Taylor principle: compatibility without a verdict

There remains the question that every estimation of this kind eventually raises: is the long-run response to inflation strong enough to be stabilising? The long-run coefficient βLT, which corrects the immediate response for inertia, comes out at about 1.23—a value above one, and thus compatible with the Taylor principle (Table 4). The imprecision is considerable: the delta-method interval [−0.264; 2.727] and the bootstrap interval [0.287; 3.964] both include unity. The bootstrap interval, however, excludes zero, which supports a positive long-run association with inflation. The implied adjustment half-life, on the order of fourteen quarters (about 3.5 years), further confirms the slowness of the conditional adjustment. The conclusion is deliberately measured: the Taylor principle is compatible with the estimation, but econometrically inconclusive. The point estimate points in the right direction; the data do not allow a firm conclusion.

Table 4. Long-run coefficient and the Taylor principle.

Indicator	Value	Interpretation
ρ (smoothing)	0.952	Very high inertia
φ_π (short run)	0.059	Positive response to inflation
$\beta_{LT} = \varphi_\pi/(1-\rho)$	1.232	Point estimate above 1
95% CI (delta)	[−0.264; 2.727]	Contains 1
95% CI (bootstrap)	[0.287; 3.964]	Contains 1
Half-life	14.1 quarters (3.5 years)	Very slow adjustment

Note. *CI: confidence interval. The bootstrap uses 2,000 residual replications.*

4.5. Robustness: what holds and what remains fragile

The robustness checks distinguish results that persist across specifications from those that depend on measurement or model choice.

For the long-run levels relationship, the reported residual diagnostics are augmented Dickey–Fuller (ADF) ≈ −4.96 and Phillips–Perron (PP) ≈ −5.59 (Table S10). Residual-based cointegration inference (Engle & Granger, 1987) requires appropriate critical values and specification details. The ARDL(2, 0, 1) bounds test reports F = 3.83 and p = 0.129 and does not provide conclusive evidence of a levels relationship (Pesaran et al., 2001). Separately, the restricted ECM reports $\Delta i_{t-1} = 0.333$ and $\Delta gap_t = 0.033$, both with $p < 0.001$; the error-correction term (ECT) is −0.046 (p = 0.054). The negative adjustment coefficient is consistent with convergence conditional on a valid levels relation, but does not establish that relation. Table 5 reports these results separately: Panel A concerns the long-run (cointegration) relationship and Panel B the short-run ECM dynamics; full details are in Tables S10–S11.

Table 5. ARDL analysis: long-run cointegration diagnostics and short-run ECM dynamics.

Statistic	Estimate	Reading
Panel A. Long run: levels (cointegration) relationship, 2008Q1–2025Q2, N = 70		
ARDL order (BIC)	ARDL(2, 0, 1)	Lag structure selected by BIC
Bounds F-test (Pesaran et al., 2001)	F = 3.83; p = 0.129	No conclusive evidence of a levels relationship

Statistic	Estimate	Reading
ADF on Engle–Granger residuals	−4.962	Requires cointegration-specific critical values
PP on Engle–Granger residuals	−5.591	Requires cointegration-specific critical values
Panel B. Short run: error-correction model (HAC standard errors in parentheses)		
Constant	0.046 (0.037); p = 0.216	Not significant
Δi(−1)	0.333 (0.071); p < 0.001	Short-run persistence of rate changes
Δgap	0.033 (0.005); p < 0.001	Short-run response to activity
ECT	−0.046 (0.024); p = 0.054	Slow adjustment, conditional on a valid long-run relationship

Note. *Panel A tests for a long-run relationship; Panel B describes short-run dynamics. ARDL: autoregressive distributed lag; ECM: error-correction model; ECT: error-correction term. Full details in Tables S10–S11.*

Across alternative output-gap measures, inertia remains high (ρ approximately 0.91–0.95), and the inflation and activity coefficients retain positive signs (Tables S12–S13). The Hamilton specification yields smaller responses, and the estimation samples differ (70, 59 and 63 observations), so the comparison combines filter and sample effects. The descriptive VAR responses show a delayed negative inflation response to an orthogonalised policy-rate innovation with wide uncertainty intervals (Tables S16–S18); their interpretation remains conditional on the recursive ordering.

The comparison with a simple autoregressive model finally puts the R^2 in its proper place. An AR(1) on the lagged rate alone already explains 93.2% of the variance, whereas the full Taylor rule reaches 95.3%: most of the fit comes from inertia, not from the economic content of the rule (Table 6). The gain in R^2 is modest (about 0.021), but the partial R^2 of 0.307 and the improvement in the AIC indicate that inflation and the output gap add explanatory content beyond inertia.

Table 6. AR(1) benchmark and the contribution of the economic variables.

Model	R^2	Adjusted R^2	AIC
AR(1) only: $i \sim i_{t-1}$	0.932	0.931	−61.05
AR(1) + inflation	0.945	0.944	−73.96
AR(1) + gap	0.940	0.938	−67.48
Simple Taylor	**0.953**	**0.951**	**−82.75**

Note. *Share of the Taylor R^2 already explained by the AR(1): 97.8%. Partial R^2 of inflation and the gap: 0.307.*

A final check concerns the most sensitive part of the data. The rule is re-estimated on a restricted CPI subsample covering 2021Q4–2025Q2 (15 observations, 21.4% of the main sample), which starts one quarter before the fully validated CPI window (Table 7). The inflation response remains positive and significant ($\varphi\pi \approx 0.116$; p = 0.035), in line with the benchmark, whereas the output-gap coefficient is not significant. The short window limits inference: this estimate is suggestive rather than confirmatory and does not validate the earlier CPI observations.

Table 7. Taylor rule on the restricted CPI subsample, 2021Q4–2025Q2 (OLS, Newey–West HAC errors).

Variable	Estimate	HAC SE	t	p
Constant	0.063	0.337	0.188	0.854
i_{t-1} (ρ)	0.848	0.085	9.937	< 0.001
π_t (φπ)	0.116	0.048	2.401	0.035
gap_t (φgap)	0.006	0.060	0.096	0.925
N	15			

Note. *The subsample represents 21.43% of the main estimation sample. With 15 observations, the estimates are imprecise and serve as a sensitivity check. Source: authors' calculations (Table S15).*

Table 8 summarises these robustness checks and the interpretive limit attached to each. Long-run cointegration and short-run ECM results appear in separate rows.

Table 8. Summary of robustness tests.

Test	Main result	Interpretive limit
Newey–West HAC errors (4 lags)	Benchmark significance of ρ, φπ and φgap	Robust covariance estimation does not correct endogeneity, measurement error or small-sample bias.
AR(1) benchmark	R^2 0.932 versus 0.953; 97.8% ratio	Most fit reflects inertia; inflation and the gap add explanatory content.
Alternative output gaps	Positive signs; N = 70, 59 and 63	Differences reflect both filtering and sample changes.
Restricted CPI subsample	N = 15 (2021Q4–2025Q2); φπ ≈ 0.116; p = 0.035	Short-window sensitivity check; does not validate earlier CPI observations. Table 7.
Long run: cointegration (ARDL bounds)	ARDL bounds F = 3.83; p = 0.129	No conclusive evidence of a levels relationship. Table 5, Panel A; Table S10.
Short run: ECM dynamics and adjustment	Δi(−1) = 0.333; Δgap = 0.033; both p < 0.001. ECT = −0.046; p = 0.054.	Short-run dynamics and conditional adjustment; ECT alone does not establish cointegration. Table 5, Panel B; Table S11.
Descriptive Cholesky VAR	Delayed inflation response; wide intervals	Conditional on recursive ordering; no established causal effect.
Stability diagnostics	Candidate breaks 2014Q2 and 2022Q1; sup-F above its 5% boundary; CUSUM within its 5% boundaries	Different alternatives and power; non-rejection is not proof of stability. Figures 3–4.

Note. *ARDL: autoregressive distributed lag; ECM: error-correction model; ECT: error-correction term; ADF: augmented Dickey–Fuller; PP: Phillips–Perron. Full details of each test are in the supplementary material.*

5. Discussion

This section interprets the estimated reaction function and the descriptive policy-mix configurations, then discusses policy implications and limitations.

5.1. A monetary policy of caution and gradualism

Strong rate persistence is consistent with a concern for predictability and financial stability, but these motivations are interpretations rather than separately identified mechanisms. As Rudebusch (2002) emphasises, persistent omitted determinants can also generate apparent smoothing. The AR(1) comparison supports a narrower statistical conclusion: most of the model fit reflects persistence, while inflation and the output gap add explanatory content (Tables 6 and 9).

Jamal-Eddine and Qachar (2026a) also report strong policy-rate persistence, together with evidence of parameter instability around the recent crises. Their analysis covers the same period, so the comparison provides closely related evidence rather than an independent replication. Unlike this companion study and Jamal-Eddine and Qachar (2026b), the present article combines the reaction-function estimates with an annual descriptive policy-mix matrix across three crises.

Table 9. Summary of results and their economic significance.

Empirical result	Status	Economic reading	Policy-mix implication
Strong policy-rate inertia	Robust	Consistent with gradualism, but also with persistent omitted determinants.	High R^2 = smoothing effect, not full explanatory power.
Positive response to inflation	Suggestive, cautious	Direction consistent with an anti-inflationary function.	CPI only partly validated before 2022: do not over-conclude.
Positive response to the gap	Positive across reported specifications	Activity is positively associated with the rate; magnitude and sample vary.	The decision is not reducible to a mechanical rule.
2022Q1 break	Candidate break	Consistent with a shift in the fitted relationship; not a separately identified policy regime.	Articulation becomes more delicate in inflationary periods.
Taylor principle	Inconclusive	Point estimate compatible, but too imprecise.	Reading grid rather than normative verdict.

Source. *Authors' synthesis.*

5.2. Observed policy configurations and the limits of coordination inference

The annual configurations show episodes in which the monetary and fiscal stances move in similar directions. Such concordance is compatible with a shared response to a common shock, separate decisions under similar constraints, or deliberate coordination. The descriptive matrix cannot distinguish these mechanisms. The evidence therefore supports a context-dependent configuration of the policy mix, without establishing either the existence or the absence of an institutional coordination framework.

Jamal-Eddine and Qachar (2026b) find no systematic direct interest-rate response to annual fiscal-stance indicators. This result reinforces the need to distinguish observed monetary–fiscal configurations from institutional coordination, which may operate through other instruments or arrangements.

Table 10 compares the economic context of the three episodes. The financial crisis calls for support against weakening external demand; the pandemic places liquidity and continuity of

activity at the centre of the response; and the inflationary episode raises the importance of price stability. These comparisons describe how the observed policy combination varies with the shock.

Table 10. Economic reading of the policy mix by crisis episode.

Episode	Dominant configuration	Economic reading	Lesson
2008–2009	2008: double restriction; 2009: expansionary concordance	Response to an external shock and weakening demand.	Support follows an initial tightening; cushioning effects are not identified here.
2020–2021	2020: exceptional support; 2021: monetary support, onset of fiscal consolidation	Priority to liquidity, activity and economic continuity.	Monetary support persists while fiscal policy begins to consolidate in 2021.
2022–2023	Tightening / double restriction	Inflation becomes the dominant risk; credibility gains weight.	Articulation grows more complex: disinflation, fiscal framing and communication must align.

Source. *Authors' synthesis, based on Figure 2 and the estimation results.*

5.3. 2022–2023: the moment when the policy mix comes under strain

The tensions in the policy mix surface above all when inflation becomes the dominant risk. So long as prices remain subdued, monetary support and fiscal support can coexist smoothly. As soon as inflation imposes itself as the priority, by contrast, objectives come into competition: disinflating without breaking activity, preserving fiscal sustainability, defending credibility and managing communication all become harder to satisfy simultaneously. The 2022–2023 sequence, identified by the breaks and by the absorption of the inflationary signal, is the moment when these tensions surface most clearly.

Imported inflation creates a difficult policy trade-off: domestic interest rates cannot directly reverse an international energy or food-price increase, while tighter financial conditions can restrain domestic demand and second-round effects. The present estimates do not measure expectations or identify these channels separately. Ouchchikh et al. (2025) provide a relevant Moroccan comparison by documenting uneven transmission across channels; Hammou Ou Ali (2025) reports a delayed housing-price response to monetary tightening. These findings make transmission lags a plausible consideration in policy design.

The first practical implication concerns the composition and communication of the policy mix during a supply-driven inflation episode. Temporary, targeted household support can be assessed alongside its financing and withdrawal conditions, so that distributional protection does not rely automatically on broad demand stimulus. A shared assessment of inflation risks, fiscal costs and debt-service sensitivity can improve consistency while preserving each institution's mandate. These are policy implications of the trade-offs discussed here, not causal effects estimated by the annual matrix.

A second implication concerns communication and financial monitoring. BAM can explain the conditions under which it adjusts or maintains the rate, distinguishing imported price pressures from signs of persistence, while monitoring lending rates, credit availability and financial vulnerabilities. The comparison with Waheed et al. (2024) makes omitted financial conditions especially relevant.

A third implication concerns debt sustainability and the data infrastructure. Liu and Mustapher (2025) illustrate, for South Africa, how a monetary–fiscal regime that contains inflation can also increase debt-service pressures. For Morocco, this motivates joint scenario analysis rather than direct transfer of their quantitative results. Publishing consistent historical series, documented national-accounts bridges and real-time data vintages would also make reaction-function evidence more useful for policy. These improvements matter because statistical precision cannot compensate for uncertain measurement.

5.4. Limitations and research agenda

The principal limitation concerns the incomplete validation of pre-2022 CPI observations in the working dataset. Year-on-year inflation uses both the current index and its value four quarters earlier, so verification of recent index levels alone does not validate every associated inflation observation.

Measurement error can affect the inflation slope, its standard error, the crisis interactions and the inferred long-run response; its bias need not be a simple attenuation when errors correlate with regressors or regimes. The 15-observation sensitivity exercise (Table 7) is too short to resolve this issue. Accordingly, the inflation-response and regime-change findings remain provisional pending an observation-by-observation reconciliation with the official series.

Further limitations concern the mechanical GDP splice, filtering choices and unequal samples in the alternative-gap checks. The estimated output gap is an ex post statistical measure, not a reconstruction of the information available to BAM at each decision date. Contemporaneous inflation and activity may be endogenous, and HAC standard errors do not address this problem. The annual matrix does not identify coordination, and the Cholesky VAR remains conditional on its ordering.

Future work should archive exact data vintages and transformation code, compare gap measures on a common sample, and incorporate expectations, energy prices, exchange rates, credit conditions and external interest rates. A joint monetary–fiscal model or a credible identification strategy is necessary before drawing causal policy-mix conclusions.

6. Conclusion

The evidence is more consistent with a sequence of partly autonomous, shock-specific responses than with an identifiable coordination mechanism, although the data cannot rule out coordination through channels not observed here. The three crises coincide with different observed monetary–fiscal configurations, and the estimated monetary reaction function displays substantial persistence and specification-dependent responses to inflation and activity.

Three conclusions follow. Policy-rate inertia is pronounced, but its interpretation as deliberate gradualism remains conditional. The inflation and output-gap coefficients are positive in the benchmark specification, while the inflation association weakens with crisis controls and remains subject to the CPI limitation. Finally, the long-run inflation coefficient is compatible with the Taylor principle, but its uncertainty prevents a definitive conclusion.

For other emerging economies, the Moroccan case illustrates the importance of distinguishing the estimation of a reaction function from the identification of policy transmission and institutional coordination. Recent research on Morocco, the BRICS and South Africa reinforces the relevance

of financial channels and fiscal conditions, while also showing why findings from different empirical designs cannot be treated as interchangeable.

For policy, the analysis motivates clearer communication of conditional decisions, better documentation of cyclical data and explicit assessment of fiscal–monetary trade-offs. A Taylor-type rule remains informative as a conditional summary of observed behaviour, provided its measurement and identification limits remain visible.

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Supplementary material. The companion file contains detailed coefficients, diagnostic tests, alternative output gaps, descriptive VAR results and policy-mix configurations.