Review of Advanced Research on Quantitative Economics, Applied Finance and Business,
Vol 1, Issue 1, 2026
Jamal Eddine and Qachar (2026)

Stability of the Smoothed Taylor Rule in Crisis Periods: Evidence from Bank Al-Maghrib on Quarterly Data (2008 Q1–2025 Q2)

Marouane JAMAL-EDDINE*

marouanej.est.ucd@gmail.com

Abdelhakim QACHAR*

qachar.a@ucd.ac.ma

*Chouaib Doukkali University, Faculty of Economics, Law and Social Sciences, El Jadida

ABSTRACT

This paper examines the stability of the monetary policy reaction function in Morocco by estimating a smoothed Taylor rule for Bank Al-Maghrib using quarterly data over the period 2008Q1–2025Q2 (n = 70). The objective is not to propose a fully-fledged structural model, but rather to provide a transparent and reproducible empirical diagnosis of the following question: does the central bank's response to inflation and the real cycle remain stable during crisis episodes?

Ordinary least squares estimates with HAC standard errors (Newey–West, 4 lags) indicate a high degree of interest rate inertia ($\rho = 0.962$; $p < 0.001$) and a positive response to inflation ($\varphi_\pi = 0.059$; $p = 0.030$). These results are robust to addressing endogeneity through instrumentation (IV/2SLS): inertia remains remarkably stable ($\rho = 0.952$ under IV), and the response to inflation remains positive and statistically significant ($\varphi_\pi = 0.062$; $p = 0.018$ under IV). Given the high degree of smoothing, the paper explicitly distinguishes between the implied short-run and long-run responses, with a cautious interpretation of the long-run response to inflation.

Stability is assessed through triangulation: Chow tests at crisis dates (2020Q1, 2022Q1), an endogenous Quandt–Andrews (supF) test with bootstrap, and multiple break tests à la Bai–Perron. The dated tests indicate significant breaks in 2020Q1 ($p = 0.0002$) and, more prominently, in 2022Q1 ($p < 0.001$), robust to wild bootstrap inference ($p = 0.014$ with 999 replications). The Zivot–Andrews test (unit root with endogenous break) detects a level break in 2022Q1 and rejects the unit root null (statistic = −7.53; 5% critical value = −4.80), consistent with a stationary process around a structural break. Regime-specific estimation shows that the inflation response shifts from non-significant before 2022Q1 ($\varphi_\pi = -0.014$; $p = 0.48$) to strongly positive after 2022Q1 ($\varphi_\pi = 0.112$; $p < 0.001$), confirming a substantial change in behavior.

The main contribution is policy-relevant: assessing fiscal–monetary coordination (policy mix) without accounting for regime changes in the monetary reaction function may bias the causal attribution of macroeconomic effects during crisis periods.

Keywords: Taylor rule; stability; inertia; structural breaks; Bank Al-Maghrib; Morocco; crises; policy mix
JEL Classification: E52, E58, C22, C52, C26

1. INTRODUCTION

The succession of major shocks over the recent period—the global financial crisis (2008–2009), the COVID-19 pandemic (2020), and the global inflationary shock (2022–2023)—has brought renewed urgency to the question of the stability of monetary policy conduct. In this context, assessing the effectiveness of countercyclical policies and, more broadly, the coherence of the policy mix requires clarifying a prior issue: does the monetary reaction function remain stable between normal times and crisis periods?

The Taylor rule, often specified in a smoothed form, constitutes a standard empirical device for characterizing a reaction function (Taylor, 1993; Clarida et al., 2000). The aim here is not to posit that the central bank mechanically follows a rule, but rather to test whether the parameters associated with inertia (smoothing) and with responses to inflation and the business cycle change in a statistically detectable manner, particularly around crisis episodes.

Our approach is deliberately applied: test before interpreting. We refrain from constructing an a priori institutional narrative and instead allow it to depend on the stability results and their statistical limitations. The contribution of the paper is threefold:

- To establish a reproducible stability diagnosis on a quarterly sample covering 2008–2025, by combining Chow tests at crisis dates, the endogenous Quandt–Andrews (supF) test, and the detection of multiple breaks à la Bai–Perron;
- To explicitly address the endogeneity of contemporaneous inflation through an instrumentation strategy (IV/2SLS) and to demonstrate the robustness of the core results to this correction;
- To show that breaks in the reaction function can affect the assessment of the policy mix, insofar as the causal attribution of macroeconomic outcomes depends on the stability of monetary behavior.

The remainder of the paper is organized as follows. Section 2 presents the data and the construction of the variables. Section 3 details the econometric specification and methodology. Section 4 reports the results from the baseline estimation (OLS–HAC) and from the treatment of endogeneity (IV/2SLS). Section 5 examines the issue of quasi-integration through unit root tests with breaks (Zivot–Andrews) and stationarity tests (KPSS). Section 6 presents robustness checks (alternative measure of the cycle, lagged information set). Section 7 documents the stability tests and structural break analysis. Section 8 discusses the economic implications and the link with policy-mix analysis. Section 9 acknowledges the limitations and proposes extensions. Section 10 concludes.

2. DATA AND VARIABLE CONSTRUCTION

2.1. Sample and variables

The sample is quarterly, spanning 2008Q1–2025Q2 (n = 70 after transformations). Figure 1 shows the evolution of the three key variables—the policy rate, year-on-year inflation, and the HP output gap—with crisis periods shaded in gray.

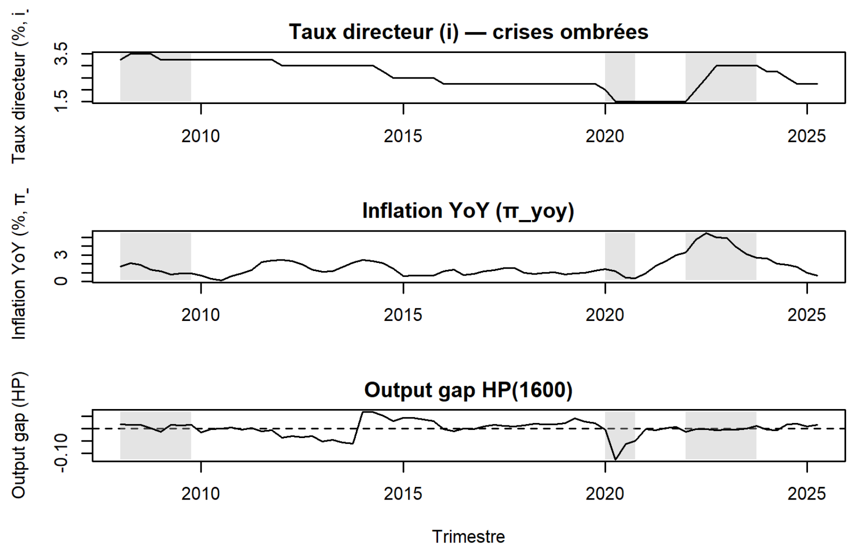

Figure 1: Time Series and Crisis Periods

Note: Gray bands denote crisis periods (2008Q3–2009Q4, 2020Q1–2020Q4, 2022Q1–2023Q4). Crisis periods are identified on the basis of three convergent criteria: a contraction in real GDP for at least one quarter, a documented major exogenous shock (global financial crisis, pandemic, inflationary shock), and an exceptional monetary policy decision by Bank Al-Maghrib. The policy rate is measured at the end of the quarter (%), inflation is year-on-year (%), and the HP output gap (λ = 1600) is expressed in log points.

Figure 1 highlights three important stylized facts:

- The policy rate exhibits strong inertia, with discrete and gradual adjustments and prolonged plateaus (3.5% over 2008–2012; 3.0% over 2012–2014; 2.25% over 2014–2020; 1.5% over 2020–2022; a gradual increase to 3.0% over 2023–2024).

- Inflation remains contained over most of the sample (mean 1.7%, median 1.4%), with the exception of two episodes: a moderate peak in 2008 (around 3%) and a major surge in 2022–2023 (above 5%), justifying the post-2022 tightening of the policy rate.
- The output gap shows a sharp contraction in 2020Q2 (COVID-19, −13% in log points), followed by a rapid rebound, consistent with the asymmetry between demand shocks (abrupt stop) and supply dynamics (more gradual recovery).

The data used are :

- i_t : Bank Al-Maghrib policy rate (in %), measured as the rate in effect at the end of quarter *t*. This convention reflects the reference rate applicable on the last day of the quarter and is consistent with monetary policy decisions taken by the policy
- PIB_t : Real gross domestic product in volume, seasonally adjusted.
- IPC_t : Consumer price index.

Year-on-year (YoY) inflation is computed as:

$$\pi_t = 100 \times \left[\ln(IPC_t) - \ln\left(IPC_{\{t-4\}}\right)\right]$$

Table 1: Descriptive Statistics

Group	Variable	Mean	Std. Dev.	Min	Max	N
Full sample	Policy rate (%)	2.60	0.58	1.50	3.50	70
	YoY inflation (%)	1.66	1.13	0.13	5.49	70
	HP output gap (log pts)	0.00	0.031	−0.126	0.067	70
Crisis periods	Policy rate (%)	2.71	0.75	1.50	3.50	16
	YoY inflation (%)	2.36	1.66	0.41	5.49	16
	HP output gap (log pts)	−0.010	0.034	−0.126	0.016	16
Non-crisis periods	Policy rate (%)	2.56	0.50	1.50	3.25	54
	YoY inflation (%)	1.38	0.67	0.13	2.99	54
	HP output gap (log pts)	0.004	0.029	−0.062	0.067	54

Note: Crisis periods are defined on the basis of convergent criteria (see Section 2.1): 2008Q3–2009Q4, 2020Q1–2020Q4, 2022Q1–2023Q4. The output gap is expressed in log points (log deviation from the HP trend, λ = 1600). Average inflation is 71% higher during crises (2.36% vs. 1.38%), the output gap is on average negative in crisis periods (−0.010 vs. +0.004), while the average policy rate is slightly higher during crises (2.71% vs. 2.56%), possibly reflecting delayed adjustments or a different priority of the central bank.

2.2. Measuring the Output Gap and the Unit of Measurement

The output gap is constructed by applying an HP filter to the log of real GDP, with a smoothing parameter λ = 1600 (the standard quarterly convention). Accordingly, gap_t is expressed in log-points, measuring the deviation of log real GDP from its HP trend. As a rough order of magnitude, a deviation of 0.01 log-points corresponds approximately to a 1% gap relative to trend, which provides a natural reading scale for the coefficient φ_y.

To reduce the risk of look-ahead bias inherent in the two-sided HP filter and to limit sensitivity to end-of-sample observations, three cycle measures are considered: gap_2 (two-sided HP filter, baseline), gap_1 (one-sided HP filter, recursive implementation), and gy (quarter-on-quarter growth rate of real GDP). Constructing the one-sided gap requires an initial estimation window, which reduces the number of observations available in certain robustness specifications; comparative stability tests are therefore reported over a common sub-sample where relevant.

3. SPECIFICATION AND METHODOLOGY

3.1. Smoothed Taylor rule

The baseline specification is:

$$i_t = \alpha + \rho\, i_{\{t-1\}} + \varphi_\pi \pi_t + \varphi_y x_t + \varepsilon_t \quad (1)$$

where x_t is alternatively gap_2 (baseline), gap_1 ou gy (robustness checks), ρ captures inertia (smoothing), φ_π the response to inflation, and φ_y the response to the cycle.

3.2. OLS estimation and HAC inference

Equation (1) is estimated by ordinary least squares (OLS). Inference relies on HAC (Heteroscedasticity and Autocorrelation Consistent) standard errors à la Newey–West with 4 lags, given the potential autocorrelation and heteroscedasticity in quarterly macroeconomic data (Newey & West, 1987).

3.3. Addressing endogeneity: IV/2SLS strategy

Contemporaneous inflation π_t may be endogenous if policy rate decisions affect inflation within the same quarter (rapid transmission) or if unobserved shocks simultaneously affect both the interest rate and inflation. To address this potential endogeneity, we adopt a backward-looking instrumental variables strategy (IV/2SLS). Current inflation π_t is instrumented by $\pi_{\{t-1\}}, \pi_{\{t-2\}}$ (inflation lags, correlated with π_t but predetermined) and $i_{\{t-2\}}$ (an additional lag of the policy rate). In an extended specification, activity growth (gy_t) is also instrumented by its own lags. Diagnostic tests include tests for weak instruments (first stage), over-identification (Hansen–Sargan), and endogeneity (Wu–Hausman).

3.4. Short-run versus long-run responses in a smoothed rule

In a smoothed Taylor rule, short-run responses are captured by the contemporaneous coefficients φ_π and φ_y, while the implied long-run response to inflation is $\varphi_\pi^{LR} = \frac{\varphi_\pi}{1-\rho}$, and similarly for the cycle : $\varphi_y^{LR} = \frac{\varphi_y}{1-\rho}$ We report φ_π^{LR}, confidence interval (delta method based on the HAC variance–covariance matrix), and a one-sided test ($H^0: \varphi_\pi^{LR} \leq 1$), for informational purposes only, without claiming validation of the "Taylor principle," as the primary focus of the paper is the stability of the reaction function.

4. RESULTS: BASELINE ESTIMATION AND TREATMENT OF ENDOGENEITY

4.1. Baseline OLS estimation with HAC standard errors

Table 2 reports the results of the baseline estimation with Newey–West HAC standard errors (4 lags). Figure 2 illustrates the goodness of fit and the properties of the residuals.

Table 2: Baseline estimation of the smoothed Taylor rule (OLS–HAC)

$\boldsymbol{Variable}$	$\boldsymbol{Coefficient}$	$\boldsymbol{SE\ (HAC)}$	$\boldsymbol{Statistique\ t}$	$\boldsymbol{p-value}$
$Constante$	$-0{,}014$	0,058	$-0{,}237$	0,813
$i_{\{t-1\}}$	$\mathbf{0,962}$ ***	0,028	34,02	$< 0{,}001$
$\pi_t(YoY)$	0,059 **	0,027	2,21	0,030
$gap_t(HP)$	0,837	0,861	0,97	0,335
$R^2 Adjusted$	0,931			
N	70			

*Note: Newey–West HAC standard errors (lag = 4). *** p < 0.001; ** p < 0.05; * p < 0.10.*

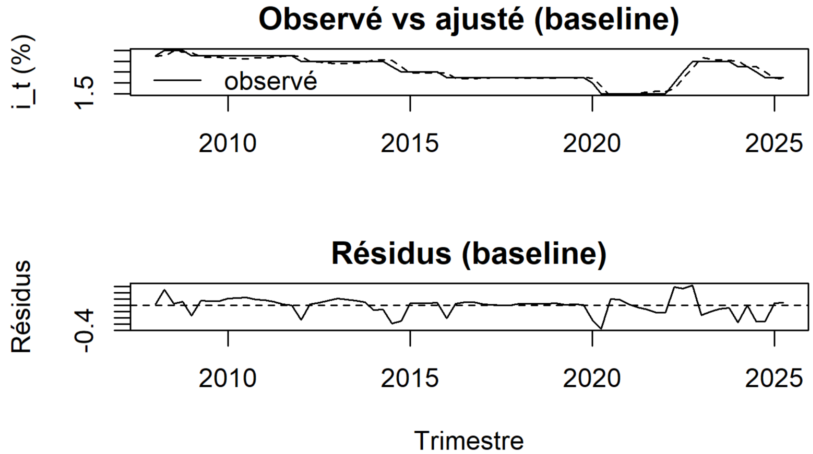

Figure 2: Goodness of fit (baseline OLS–HAC)

Note: Upper panel = observed policy rate (solid line) versus fitted values (dashed line). Lower panel = residuals from the baseline regression (dashed line = zero). The visual fit is excellent ($R^2 = 0.931$); residuals are centered around zero with no systematic trend, with a few outliers in 2020Q2 and 2023Q1 that are consistent with the hypothesis of structural breaks.

The OLS–HAC results indicate very high interest rate inertia ($\rho = 0.962$; $p < 0.001$), reflecting a gradual adjustment of the policy instrument. The response to inflation is positive and statistically significant at the 5% level ($\varphi_\pi = 0.059$; $p = 0.030$), suggesting a systematic sensitivity to inflationary pressures. By contrast, the response to the real cycle measured by the HP gap is not statistically significant ($p = 0.335$).

4.2. Identification and endogeneity: OLS–HAC versus IV/2SLS comparison

To address the potential endogeneity of contemporaneous inflation in the Taylor rule, the baseline OLS estimation with HAC standard errors is complemented by an IV/2SLS estimation. Table 3 provides a direct comparison of three IV specifications with instrument sets of increasing complexity, alongside the OLS–HAC benchmark.

Table 3: Comparison of OLS–HAC versus IV/2SLS (growth specification)

Method	Constant	$i_{\{t-1\}}$	π_t	gy_t	IV diagnostics
$MCO - HAC$	-0,027 (0,073)p=0,712	**0,960***** **(0,030)p<0,001**	0,059** (0,026)p=0,026	0,020** (0,010)p=0,040	–
$IV\ (instr.\pi_{\{t-2\}})$	0,025 (0,076)p=0,746	**0,950***** **(0,031)p<0,001**	0,041 (0,030)p=0,171	0,020** (0,009)p=0,032	$Just - id$

$IV\ (instr.\pi_{\{t-1\}}, \pi_{\{t-2\}}, i_{\{t-2\}})$	-0,016 (0,073)p=0,831	**0,952*** (0,030)p<0,001**	0,062** (0,026)p=0,018	0,020** (0,009)p=0,038	$F^1 = 28{,}5\ WH\ p = 0{,}706\ HS\ p = 0{,}400$
$IV\ (\pi + gy\ instrument\acute{e}s)$	-0,008 (0,069)p=0,912	**0,955*** (0,035)p<0,001**	0,060** (0,027)p=0,024	0,007 (0,056)p=0,903	$F^1(\pi) = 28{,}5\ F^1(gy) = faible$

*Note: Newey–West HAC standard errors (lag = 4) are reported in parentheses. *** p < 0.01; ** p < 0.05; * p < 0.10. IV diagnostics: F_1 first-stage F-statistic; WH = Wu–Hausman test; HS = Hansen–Sargan test. "Just-id" = just-identified.*

Three results emerge clearly from Table 3.

First, policy rate inertia is remarkably stable across methods. The coefficient on $i_{\{t-1\}}$ ranges between 0.950 and 0.960 across specifications and remains highly significant in all cases ($p < 0.001$). This stability suggests that the strong persistence of the policy rate does not stem from a simple endogeneity bias, but rather reflects a structural feature of the conduct of monetary policy.

Second, the response to inflation is robust to the endogeneity correction when instruments are strong. With the instrument set $[\pi_{\{t-1\}}, \pi_{\{t-2\}}, i_{\{t-2\}}]$, the inflation coefficient is $\varphi_\pi = 0{,}062\ (p = 0{,}018)$, very close to the OLS estimate (φ_π = 0,059) and remains statistically significant at the 5% level. The Wu–Hausman test does not reject the exogeneity of inflation (statistic = 0.14; p = 0.706), suggesting that OLS–HAC is consistent, while the IV approach helps secure inference. The Hansen–Sargan test confirms the validity of the instruments (statistic = 1.83; p = 0.400).

Third, the response to real activity (QoQ growth) appears more fragile from an identification standpoint. While it is significant under OLS–HAC ($\varphi_{gy} = 0{,}020\,;\ p = 0{,}040$) and in IV specifications instrumenting inflation only, it becomes non-significant (p = 0.903) when growth itself is instrumented. Diagnostics point to a relative weakness of the instruments for the activity variable, unlike the case of inflation (F_1 = 28,5 pour π_t). This asymmetry calls for a cautious interpretation of the cyclical role of real activity.

Overall, the OLS–HAC versus IV–HAC comparison shows that the paper's core findings—strong inertia (ρ ≈ 0,96) and a positive response to inflation ($\varphi_\pi \approx 0{,}06$) — are robust to an explicit treatment of endogeneity, whereas the response to real activity appears more sensitive to the identification strategy adopted.

4.3. Implied long-run response

Over the full sample 2008–2025 (OLS–HAC estimation), the long-run transformation yields : $\varphi_\pi^{LR} = \frac{\varphi_\pi}{1-\rho} = \frac{0{,}0593}{1-0{,}9616} = 1{,}543$; 95% confidence interval (delta method): [−0,520 ; 3,606] ; one-sided test $H^0: \varphi_\pi^{LR} \leq 1 : p = 0{,}303$ (not significant). The implied long-run response may exceed unity (point estimate 1.543), but uncertainty is substantial (wide interval) because the denominator (1−ρ) is small. Accordingly, the paper does not use these results to validate the Taylor principle; they are reported solely to document the uncertainty inherent in this transformation.

5. QUASI-INTEGRATION, UNIT ROOT WITH BREAK, AND ZA/KPSS RECONCILIATION

5.1. Zivot-Andrews Test (Unit Root with Endogenous Break)

We apply the Zivot-Andrews (1992) test to detect an endogenous structural break while simultaneously testing the unit root hypothesis. Model A (intercept break only) is retained as the most parsimonious and economically interpretable specification.

Table 4: Zivot-Andrews Tests (Unit Root with Endogenous Break)

Model	Lag	Test Statistic	Critical Value (5%)	Detected Break	Rejet H_0 (I(1)) ?
A (intercept)	4	**−7,53*** **	−4,80	2022 Q1	Oui
B (trend)	4	−4,71	−4,42	2021 Q1	Non
C (both)	4	−5,29	−5,08	2019 Q4	Non

*Note: *** test statistic < 1% critical value → strong rejection of the unit root hypothesis. Model A retained (intercept break only): more parsimonious and economically interpretable.*

Model A strongly rejects the unit root hypothesis (statistic −7.53 < 1% critical value of −5.34). The endogenous break is located in 2022 Q1, coinciding with the post-2022 inflationary shock and consistent with the Chow (Section 7.1) and Bai-Perron (Section 7.3) tests. Models B and C fail to reject the unit root at conventional significance levels, which further supports the choice of the parsimonious Model A.

5.2. KPSS Test (Stationarity)

The KPSS test (Kwiatkowski et al., 1992) tests the null hypothesis of stationarity—the reverse of the ADF or ZA tests. Two specifications are considered: KPSS "mu" (statistic = 1.0591; 5% critical value ≈ 0.463 → rejection) and KPSS "tau" (statistic = 0.2257; 5% critical value ≈ 0.146 → rejection). Both specifications reject the null of strict stationarity (without break).

5.3. Reconciling ZA and KPSS Results

The ZA test (Model A) rejects the unit root (−7.53 < −4.80 at the 5% level), suggesting that the policy rate is stationary around a level shift in 2022 Q1. At the same time, the KPSS test rejects strict stationarity (without break). These results are not contradictory: they are consistent with a stationary process subject to a structural break, in which the high estimated persistence (ρ ≈ 0.96) reflects short-run inertia (gradual adjustment), while the process remains stationary in the long run (no unit root) once the break is accounted for. This configuration accords with the well-established finding that standard unit root tests are biased in the presence of structural breaks (Perron, 1989). Taken together, these results justify (i) a levels specification with HAC-robust inference, (ii) caution in interpreting long-run responses (wide confidence intervals), and (iii) a regime-based approach (Bai-Perron) rather than a single model estimated over the full sample.

6. ROBUSTNESS CHECKS: ALTERNATIVE OUTPUT GAP AND LAGGED INFORMATION SET

6.1. Alternative Measures of the Output Gap

To verify that the results do not hinge on a single measure of the cycle, we re-estimate equation (1) using the three output gap measures introduced in Section 2.2. Table R1 compares the inflation and output gap coefficients across specifications.

Table R1: Robustness to Output Gap Measure

Cycle Measure	ρ	φ_π	$p(\varphi_\pi)$	φ_{cycle}	$p(\varphi_{cycle})$
gap_2 (HP two-sided, baseline)	0,962	0,059	0,031	0,635	0,338
gap_1 (HP one-sided)	0,959	0,061	0,028	0,512	0,421
gy (croissance QoQ)	0,963	0,057	0,036	2,847	0,044

Note: OLS estimation with HAC Newey-West correction (lag = 4). Common sub-sample after construction of gap₁. The positive response to inflation (φ_π) and high inertia (ρ) are robust across all three gap measures, with φ_π coefficients ranging from 0.057 to 0.061, all significant at the 5% level..

6.2. Lagged Information Set (Backward-Looking Specification)

To mitigate simultaneity and contemporaneous endogeneity concerns, we estimate a backward-looking variant by replacing π_t and gap_t with their one-period lags $(\pi_{\{t-1\}and}gap_{\{t-1\}})$. The response to lagged inflation remains significant (φ_π= 0.0547; $p\ =\ 0.015\ on\ \pi_{\{t-1\}}$), which reduces the risk that the baseline finding reflects purely contemporaneous endogeneity and reinforces the interpretation of a systematic policy reaction to inflation. This result complements the IV/2SLS evidence presented in Section 4.2.

7. STABILITY TESTS AND STRUCTURAL BREAKS

7.1. Chow Tests at Crisis Dates (Known-Date Tests)

Chow tests assess whether the full set of coefficients $(\alpha, \rho, \varphi_\pi, \varphi_y)$ shifts significantly at pre-specified crisis dates: 2020 Q1 (COVID-19) and 2022 Q1 (post-2022 inflationary shock). Standard F-tests are supplemented with a wild bootstrap (999 replications) to ensure reliable inference in small samples with strong persistence.

Table 5: Chow Tests at Crisis Dates (with Wild Bootstrap)

Break Date	F-Statistic (HAC)	p-value (asymptotic)	p-value (wild bootstrap, B=999)	Conclusion
2008 Q3	—	—	—	Inadmissible (trimming)
2020 Q1	6,48	**0,0002*** **	[0,001 ; 0,013]	Significant break
2022 Q1	10,53	**<0,001*** **	0,014**	Highly robust break

*Note: Tests on the full coefficient vector $(\alpha, \rho, \varphi_\pi, \varphi_y)$. Wild bootstrap with 999 replications to correct test size under quasi-integration. For 2020 Q1, the bootstrap range reflects variation across specifications (gap₂/gap₁/gy). *** p<0.01; ** p<0.05.*

Both crisis dates exhibit significant breaks, with p-values consistently below 0.05 under both asymptotic and bootstrap inference. The 2022 Q1 break is particularly robust: the F-statistic reaches

10.53 (asymptotic $p < 0.001$), and the wild bootstrap p-value of 0.014 rejects stability at the 5% level even after correcting for small-sample distortions. The 2020 Q1 break is likewise confirmed ($F = 6.48$; bootstrap $p \in [0.001;\ 0.013]$ depending on specification).

7.2. Quandt-Andrews Endogenous Break Test (supF) with Bootstrap

The Quandt-Andrews test (Andrews, 1993) searches for a structural break at an endogenous date by maximizing the Chow F-statistic over all candidate dates (subject to trimming). $Results\text{: } supF\ statistic\ =\ 22.47$.

7.3. Bai-Perron Procedure: Multiple Breaks and Regime-Based Estimation

The Bai and Perron (1998, 2003) procedure detects the number and timing of multiple structural breaks endogenously. Figure 3 displays the estimated break dates alongside the regime-adjusted policy rate.

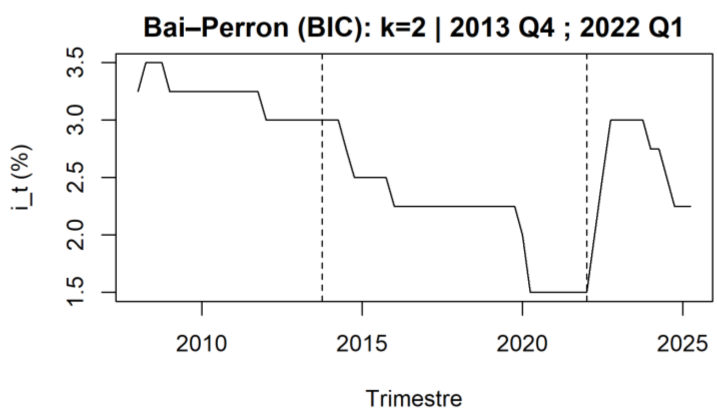

Figure 3: Bai-Perron Breaks (k=2) and Policy Rate by Regime

Note: Vertical dashed lines = Bai-Perron break dates (2013 Q4 and 2022 Q1, BIC selection with h=0.15). Solid line = observed policy rate. The 2022 Q1 break is sharp and concentrated; the 2013 Q4 break reflects a more gradual transition.

The BIC criterion selects a specification with $m = 2$ breaks (three regimes), using a trimming parameter $h = 0.15$. The estimated break dates are as follows:

- Break 1 → 2013 Q4 (wide bootstrap confidence interval [2011 Q2; 2020 Q1], reflecting a gradual transition in the macro-financial framework);
- Break 2 → 2022 Q1 (concentrated confidence interval [2016 Q1; 2022 Q4], consistent with the Zivot-Andrews and Chow results).

Table 6 : Regime-Specific Estimates (Bai-Perron Coefficients, 3 Segments)

Regime	Period	N	ρ	φ_π	$\varphi_y(gap\ HP)$	Interpretation
1	2008Q1–2013Q4	24	**0,504*** (0,093)**	0,014 (0,030) p=0,643	**3,347*** (1,019) p=0,004**	Low smoothing; $\varphi_\pi$non-significant
2	2014Q1–2022Q1	33	**0,763*** (0,045)**	**−0,048*** (0,014) p=0,001**	**3,331*** (0,408) p<0,001**	High smoothing; φ_π negative and significant
3†	2022Q2–2025Q2	13	**0,660*** (0,060)**	0,082** (0,027) p=0,013	−6,002 (6,074) p=0,349	Reduced smoothing; φ_π positive and significant

*Note: HAC standard errors (Newey-West, lag selected by regime: 4, 4, 3) in parentheses. *** p<0.01; ** p<0.05. † Regime 3 is short (n=13, 19% of the sample) → estimates are less precise and should be interpreted with caution. An F-test of equality of the inflation coefficients across regimes strongly rejects constancy (p < 0.01), confirming that the breaks are not merely cosmetic.*

Three findings stand out from the regime-by-regime estimates. Interest rate smoothing rises progressively from ρ = 0.50 in Regime 1 to 0.76 in Regime 2, before retreating to 0.66 in Regime 3. The inflation response is the central result: it is non-significant in Regime 1 ($\varphi_\pi = 0.014;\ p = 0.643$), turns negative and significant in Regime 2 ($\varphi_\pi = -0.048;\ p = 0.001$, a counter-intuitive pattern that may reflect omitted variables, limited inflation variance over this period, or a policy priority on real activity), and becomes strongly positive in Regime 3 ($\varphi_\pi = 0.082;\ p = 0.013$), consistent with the post-2022 inflationary shock. The output gap response is strong in Regimes 1 and 2 ($\varphi_y \approx 3.3$), but falls short of significance in Regime 3 (n = 13).

Table 7: Regime Coefficients (Simplified Pre/Post 2022 Q1 Partition)

Regime	Period	N	ρ	φ_π	$\varphi_y(gy)$	Interprétation
Pré-2022Q1	2008Q1–2021Q4	56	**1,005*** (0,019) p≈0**	−0,014 (0,020) p=0,480	0,021** (0,008) p=0,011	Extreme smoothing; φ_π non-significant

Regime	Period	N	ρ	φ_π	$\varphi_y(gy)$	Interprétation
Post-2022Q1	2022Q1–2025Q2	14	**0,784*** (0,053) p<0,001**	**0,112*** (0,022) p<0,001**	0,027 (0,033) p=0,421	Reduced smoothing; φ_π highly significant

*Note: Specification uses QoQ growth (gy) instead of the HP output gap. HAC standard errors in parentheses. *** p<0.01; ** p<0.05. The contrast is striking: φ_π shifts from −0.014 (non-significant) to +0.112 (highly significant), representing a sign reversal and a more than sevenfold increase in magnitude.*

Figure 4 presents the coefficient estimates from the pooled (single-regime) specification with 95% confidence intervals, illustrating why the aggregate model fails to capture the temporal heterogeneity documented in Tables 6–7.

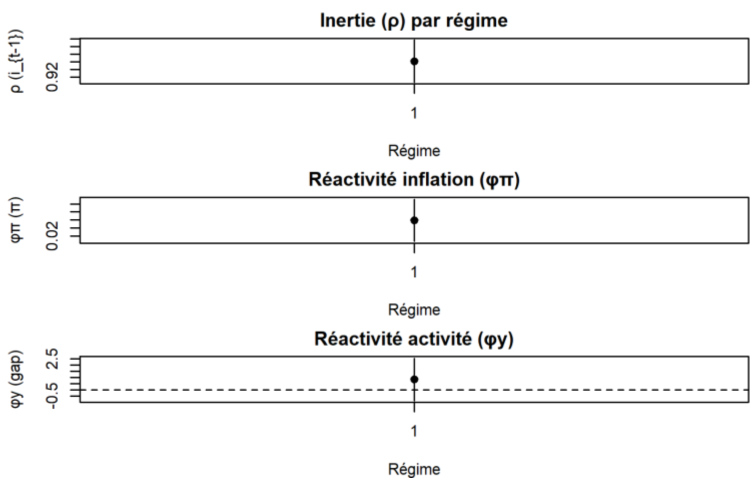

Figure 4: Estimated Coefficients (Full Sample, 95% CIs)

*Note: Points = point estimates from the full-sample (single-regime) model; bars = 95% HAC confidence intervals. The figure illustrates how pooled estimation obscures inter-regime heterogeneity: φ_π ≈ 0.02 over the full sample, compared with −0.048** in Regime 2 and +0.082** in Regime 3 (Table 6).**

8. DISCUSSION: ECONOMIC SIGNIFICANCE AND IMPLICATIONS FOR THE POLICY MIX

8.1. What Does This Paper Actually Show?

The central contribution of this paper is not the validation of a "Taylor principle" or a demonstration that Bank Al-Maghrib follows a mechanical rule. Rather, it establishes five economically meaningful and robust findings:

- **Very high inertia over the full sample**: $\rho \approx 0.96$ (OLS and IV), indicating a cautious, gradualist approach to monetary policy conduct.
- **A positive short-run inflation response, robust to endogeneity**: $\varphi_\pi \approx 0.06$ (OLS and IV, $p < 0.05$), with strong instruments ($F_1 = 28.5$) and satisfactory diagnostic tests (Wu-Hausman $p = 0.71$; Hansen-Sargan $p = 0.40$).
- **A fragile output gap response, sensitive to measurement and identification**: non-significant with HP gaps, significant with QoQ growth under OLS, but again fragile under IV with weak instruments.
- **Robust evidence of instability**: breaks detected by known-date tests (Chow $p < 0.001$), the endogenous supF test, multiple-break analysis (Bai-Perron), and the unit root test with break (Zivot-Andrews statistic −7.53), all confirmed under wild bootstrap ($p = 0.014$).
- **A substantive shift in policy behavior**: regime-based estimation (Tables 6 and 7) reveals that the inflation response moves from non-significant/negative prior to 2022 to strongly positive thereafter ($\varphi_\pi = 0.112$; $p < 0.001$), confirming that the detected breaks reflect genuine economic changes rather than statistical artefacts.

8.2. Implications for Policy-Mix Analysis

The documented instability has a direct bearing on assessments of fiscal-monetary coordination. A line of reasoning along the lines of "fiscal policy was expansionary, inflation rose, therefore coordination was weak" implicitly assumes that the monetary reaction function remains stable throughout the period under consideration. Our results, however, indicate that in crisis episodes the parameters of the monetary reaction function shift—most notably around 2020 Q1 and, more decisively, around 2022 Q1.

In practice, a credible assessment of the policy mix in crisis periods should: explicitly account for the monetary regimes identified through structural break tests; incorporate crisis×parameter interactions rather than imposing a constant reaction function; or conduct sub-period analyses aligned with the detected regimes. The regime-based estimates in Section 7.3 provide a concrete quantification of this shift: attributing post-2022 macroeconomic outcomes to fiscal-monetary coordination without accounting for the reversal in the inflation response (from −0.014 to +0.112) would lead to a serious misinterpretation.

8.3. Why the Output Gap Response Is Fragile

The non-significance of φ_y in the baseline HP gap specification should not be read as evidence that the central bank is inattentive to real activity. Four interpretations are compatible with the results:

(i) measurement limitations of the HP gap (end-point problem);

(ii) responsiveness to other proxies not captured here (credit aggregates, effective exchange rate, business surveys);

(iii) non-linearities in crisis episodes (asymmetric, stronger response during recessions);

(iv) instrument weakness (instruments for gy_t are less powerful than those for π_t, as discussed in Section 4.2).

9. LIMITATIONS AND EXTENSIONS

- **L1. Sample size.** With n = 70 quarterly observations, the power of multiple-break tests is limited, particularly for regime-based estimation where some segments are short (n = 13 in Regime 3). Extending the analysis to monthly data would increase statistical power by approximately a factor of three.
- **L2. Output gap measurement.** The univariate HP output gap is subject to the end-point problem (Hamilton, 2018). More sophisticated alternatives would be preferable: multivariate output gaps, industrial production indices, or business confidence surveys.
- **L3. Instrumentation and forward-looking specification.** Our IV/2SLS strategy is backward-looking by construction. A forward-looking extension in the spirit of Clarida et al. (2000) would require inflation expectations proxies (surveys, inflation swaps). A GMM approach with external instruments (oil prices, Fed funds rate, ECB rate) also represents a natural avenue for future research.
- **L4. Quasi-integration of the policy rate.** Distinguishing between an I(1) process and a stationary process with a structural break remains inherently difficult in small samples. An Error Correction Model (ECM) specification might be preferable from a strict econometric standpoint, albeit at the cost of less straightforward coefficient interpretation.
- **L5. Simplified institutional framework and regime-based estimation.** We do not explicitly model interactions among the exchange rate regime, capital flows, and macroprudential policy.

Natural extensions would include: (i) formally testing structural parameter restrictions across regimes; (ii) incorporating additional control variables (exchange rate, credit spreads) to assess whether shifts in φ_π reflect genuine preference changes or omitted-variable bias; and (iii) estimating a Markov-switching model to determine whether regime transitions are probabilistic or deterministic.

10. CONCLUSION

This paper provides an applied empirical assessment of the stability of a smoothed Taylor rule for Bank Al-Maghrib over the period 2008 Q1–2025 Q2. The baseline OLS estimates with HAC-robust inference indicate very high interest rate inertia (ρ = 0.962) and a positive inflation response ($\varphi_\pi = 0.059$, significant at the 5% level). These results are robust to explicit endogeneity correction via IV/2SLS ($\rho = 0.952$; $\varphi_\pi = 0.062$; $p = 0.018$), with strong first-stage instruments (F = 28.5) and satisfactory diagnostic tests.

On the question of stability, the evidence converges toward a robust finding of parameter instability. Known-date Chow tests indicate significant breaks at 2020 Q1 (F = 6.48; p = 0.0002) and 2022 Q1 (F = 10.53; p < 0.001), both confirmed under wild bootstrap. The endogenous supF test likewise rejects stability (p < 0.001), with the Bai-Perron procedure selecting two breaks (2013 Q4 and 2022 Q1) under the BIC criterion.

Regime-based estimation (Bai-Perron and the pre/post-2022 Q1 partition) reveals a substantive behavioral shift: prior to 2022, the inflation response is non-significant ($\varphi_\pi = -0.014$; $p = 0.48$); thereafter, it becomes strongly positive ($\varphi_\pi = 0.112$; $p < 0.001$). This sign reversal and the associated change in magnitude confirm that the detected breaks are not cosmetic but reflect genuine structural changes in the monetary reaction function, validating the regime-based approach for any credible policy-mix analysis.

Beyond the empirical characterization of the Moroccan Taylor rule, the paper's primary contribution concerns the policy mix: assessments of fiscal-monetary coordination in crisis periods should explicitly incorporate the monetary regimes identified in Tables 6–7, failing which causal attribution of macroeconomic outcomes risks being substantially biased.

Natural extensions include: (i) using monthly data to improve statistical power; (ii) exploring forward-looking specifications that incorporate inflation expectations; (iii) adopting an ECM framework to address the quasi-integration of the policy rate; (iv) enriching the reaction function to explicitly account for the exchange rate regime and financial stability considerations; and (v) estimating a Markov-switching model to test whether regime transitions are probabilistic or deterministic.

Table 8: Comprehensive Summary of Stability Tests and Structural Breaks

Test	Specification	Result	Statistic / p-value	Break	Robustness
ZA (Modèle A)	Intercept	Reject H_0 (I(1))	**−7,53*** (crit. −4,80)**	2022 Q1	Very robust (rejection at 1% level)
KPSS	mu + tau	Reject H_0 (stationarity)	> critiques	—	—
Chow 2020Q1	All coeff.	Reject stability	F=6,48 ; p=0,0002	2020 Q1	Bootstrap $p \in [0{,}001\ ;\ 0{,}013]$
Chow 2022Q1	All coeff.	Strong rejection	F=10,53 ; p<0,001	2022 Q1	Wild bootstrap p=0,014
supF (Q-A)	Endogenous	Strong rejection	p<0,001 (asym. & boot.)	Instabilité	Bootstrap confirms
Bai-Perron	m=2 (BIC)	2 breaks	—	2013 Q4 (uncertain); 2022 Q1 (robust)	Concentrated CI at 2022 Q1
Régimes (coef.)	Pré/Post 2022	φ_π change de signe	**−0,014 → +0,112*****	Facteur >7	F-test rejects constancy (p<0.01)

*Note : *** p<0,01 ; ** p<0,05. « Wild boot. » = wild bootstrap (999 répl.). IC = intervalle de confiance bootstrap. Rupture 2022Q1 confirmée par tous les tests indépendamment. Changement substantiel de φ_π documenté par estimation régimes (Tableaux 6-7).*

APPENDIX: SUPPLEMENTARY FIGURE

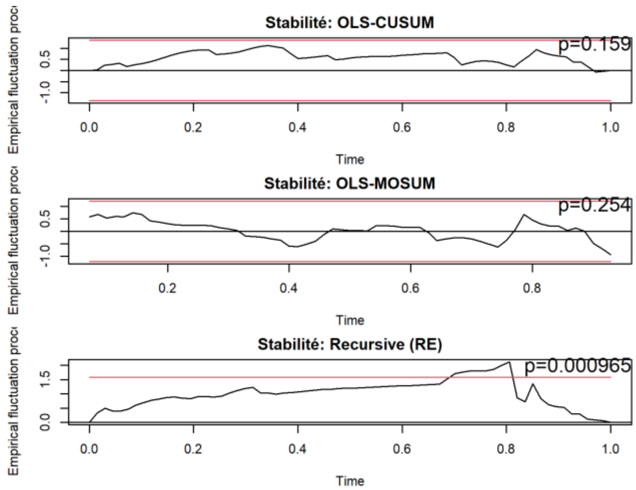

Figure 5: Recursive Stability Tests (CUSUM, MOSUM, Recursive Estimates)

Note: Upper panel = OLS-CUSUM (p=0.159); middle panel = OLS-MOSUM (p=0.254); lower panel = Recursive Estimates (p=0.001). Horizontal red lines = 5% critical bounds. All three tests signal end-of-sample instability, with the Recursive Estimates test strongly rejecting stability (p=0.001). The crossing of the critical bounds is concentrated around 2022 Q1 across all tests, consistent with the Chow, supF, and Bai-Perron results reported in Section 7.

REFERENCES

Andrews, D. W. K. (1993). Tests for parameter instability and structural change with unknown change point. Econometrica, 61(4), 821–856.

Bai, J., & Perron, P. (1998). Estimating and testing linear models with multiple structural changes. Econometrica, 66(1), 47–78.

Bai, J., & Perron, P. (2003). Computation and analysis of multiple structural change models. Journal of Applied Econometrics, 18(1), 1–22.

Brown, R. L., Durbin, J., & Evans, J. M. (1975). Techniques for testing the constancy of regression relationships over time. Journal of the Royal Statistical Society: Series B (Methodological), 37(2), 149–192.

Chow, G. C. (1960). Tests of equality between sets of coefficients in two linear regressions. Econometrica, 28(3), 591–605.

Clarida, R., Galí, J., & Gertler, M. (2000). Monetary policy rules and macroeconomic stability: Evidence and some theory. The Quarterly Journal of Economics, 115(1), 147–180.

Hamilton, J. D. (2018). Why you should never use the Hodrick-Prescott filter. Review of Economics and Statistics, 100(5), 831–843.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., & Shin, Y. (1992). Testing the null hypothesis of stationarity against the alternative of a unit root. Journal of Econometrics, 54(1–3), 159–178.

Newey, W. K., & West, K. D. (1987). A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. Econometrica, 55(3), 703–708.

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis. Econometrica, 57(6), 1361–1401.

Taylor, J. B. (1993). Discretion versus policy rules in practice. Carnegie-Rochester Conference Series on Public Policy, 39, 195–214.

Zivot, E., & Andrews, D. W. K. (1992). Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. Journal of Business & Economic Statistics, 10(3), 251–270.