



From r^* to monetary policy stance: Evidence from Poland

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Abstract

This paper estimates Poland's time-varying natural real interest rate and natural-rate gap over 2010Q1–2025Q1 and examines how a natural-rate-gap innovation propagates in a Structural Vector Error Correction (SVEC) framework. The measurement stage uses a Laubach-Williams-type state-space model disciplined by an ex ante real policy rate constructed from survey-based inflation expectations. The transmission stage embeds the estimated gap in a cointegrated macro-financial system covering activity, inflation, the real effective exchange rate, the term spread, debt-service costs and the natural-rate gap, with domestic and external controls. The estimated natural rate declines around the pandemic period and rises sharply during the post-pandemic tightening episode. Robustness analysis based on alternative expectations measures, GDP/GVA activity indicators, moderate calibration variants and one-sided filtering preserves the main cyclical profile, while highly flexible natural-rate specifications behave as diagnostic stress tests. SVEC results indicate bounded and weakly estimated transmission: statistically informative responses are sparse, CPI provides the clearest limited nominal evidence, and forecast-error-variance contributions remain small for most macro-financial variables. The natural-rate gap is therefore informative for monetary-stance assessment, but not a standalone explanation of Polish macroeconomic fluctuations.

Keywords

- natural real interest rate
- natural-rate gap
- monetary-policy stance
- inflation expectations
- SVEC

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Introduction

The natural real interest rate is a central but unobservable benchmark in monetary economics. In the Wicksellian tradition, it is the real rate consistent with price stability and macroeconomic equilibrium and not a directly observed market yield (Wicksell, 1936). In contemporary policy analysis, the concept matters because the same nominal policy rate can imply a restrictive or accommodative stance depending on inflation expectations and the level of the natural rate.

For a small open economy such as Poland, this issue is particularly relevant. Domestic monetary conditions interact with external financial cycles, exchange-rate movements, capital flows and risk premia. Assessing the monetary-policy stance only through nominal policy rates may therefore obscure the real intertemporal wedge facing households, firms and financial intermediaries. A natural-rate-gap measure, defined as the difference between the *ex ante* real policy rate and the estimated natural real interest rate, provides a compact indicator of this wedge. Woodford (2003) establishes the role of the natural-rate gap in modern monetary-policy theory by linking real-rate deviations from equilibrium to intertemporal demand pressure.

This paper has two objectives. First, it estimates Poland's natural real interest rate and natural-rate gap over 2010Q1–2025Q1 using a semi-structural state-space model. The real-rate input is constructed *ex ante*, using survey-based inflation expectations instead of realised inflation. Second, it embeds the resulting stance measure in a structural vector error-correction framework in order to examine how a natural-rate-gap innovation propagates to activity, inflation, the real effective exchange rate, the term spread, unemployment and debt-service costs.

The contribution is empirical and methodological. On the measurement side, the paper combines survey-based expectations, Kalman filtering and a broad robustness audit covering expectations, activity measures, calibration choices and end-of-sample filtering. On the transmission side, the paper treats the natural-rate gap as the structural object of interest, not only as a descriptive series. The resulting framework is especially useful for economies in which the policy-relevant real rate is difficult to infer from nominal instruments alone. Recent policy work has emphasised that stance measurement should explicitly confront uncertainty surrounding the natural rate and the *ex ante* real rate (Harrison & Nguyen, 2025).

The empirical interpretation is cautious. The natural-rate estimate is informative but not observed; the natural-rate gap is policy-relevant, but uncertain; and the SVEC evidence is conditional on the measurement stage and the short quarterly sample. The results show a sharp post-pandemic rise in the natural real rate, a strongly restrictive stance during the tightening cycle, and a transmission pattern in which CPI inflation and selected external-financial variables respond more systematically than real activity and debt-service costs.

The remainder of the paper is organised as follows: Section 1 reviews the relevant literature, Section 2 presents the data, notation and empirical strategy, Section 3 reports the results and discusses robustness evidence, including the appendix diagnostics where relevant, and Section 4 discusses policy implications and limitations. Last but not least, Section 5 concludes the entire paper.

1. Literature review

1.1. Natural real interest rate measurement

The empirical literature typically treats r^* as a latent state. Laubach and Williams (2003) provide the canonical semi-structural state-space framework in which potential output, trend growth, the output gap and the natural real rate are jointly inferred from macroeconomic observables. This approach is attractive because it imposes economic discipline while allowing the equilibrium real rate to vary over time.

International extensions of this framework show that natural-rate estimates can move together across economies and are sensitive to long-run growth, demographics and global financial conditions. Holston et al. (2017) extend the measurement problem to several advanced economies and demonstrate that the decline in equilibrium real rates is not confined to one country. Kiley (2020) stresses that the data contain limited information about low-frequency natural-rate movements, which makes uncertainty an inherent feature of the exercise. Hamilton et al. (2016) reach a similar conclusion from a longer-run equilibrium-rate perspective, emphasising that the link between trend growth and the equilibrium real rate is weaker and less stable than simple growth-based interpretations suggest. Rachel and Summers (2019) add a global perspective by linking lower neutral rates to broad saving-investment forces and fiscal off-sets in industrial economies.

Several refinements are directly relevant to the present paper. Lewis and Vazquez-Grande (2019) show that transitory shocks can materially distort natural-rate estimates when the state-space structure is overly restrictive. Orphanides and van Norden (2002) document the unreliability of real-time output-gap estimates, which motivates the explicit comparison between smoothed and one-sided filtered natural-rate paths. Buncic (2024) further emphasises the econometric fragility of natural-rate estimation, including sensitivity to signal extraction, persistence assumptions and sample endpoints.

Recent microdata-based evidence for Poland shows that firm-level business-tendency and enterprise-report information can generate output-gap measures with useful cyclical and Phillips-curve information, including quasi-real-time properties (Błażej et al., 2025). The present paper treats output-gap uncertainty as a reason for explicit filtering robustness and transparent aggregate state-space measurement.

Recent post-pandemic evidence reinforces the need for robustness analysis. Benigno et al. (2024) argue that the natural real rate may have shifted upward after the pandemic, although the magnitude and persistence of that shift remain uncertain. Del Negro et al. (2017) show that safety and liquidity premia can depress natural-rate estimates, which is especially relevant when financial-market conditions and risk premia influence small open economies. Brand, Goy and Lemke (2026) show how macro-finance information can discipline natural-rate estimates. Brand, Lisack and Mazelis (2025) summarise euro-area natural-rate estimates and stress that point estimates should be interpreted together with uncertainty bands. Juselius (2025) similarly frames r^* as a range of plausible estimates.

1.2. Inflation expectations and the ex ante real rate

A credible natural-rate-gap measure requires an ex ante real policy rate. Realised inflation is not sufficient when inflation is volatile or when survey expectations adjust with a lag. Carlson and Parkin (1975) introduced a probabilistic method for converting qualitative survey responses into quantitative inflation expectations, which remains useful when direct numerical expectations are unavailable.

The role of inflation expectations in natural-rate measurement has received renewed attention. Lopez-Salido et al. (2020) show that alternative expectations measures can materially affect the inferred equilibrium real rate. Coibion and Gorodnichenko (2025) review the evidence on expectation formation and emphasise that survey expectations often differ from model-consistent or market-implied measures. This motivates the paper's comparison of the baseline Carlson-Parkin expectations series with ordered-survey and model-implied alternatives.

The policy interpretation of the natural-rate gap is grounded in New Keynesian monetary economics. Clarida et al. (1999) link stabilisation policy to the interaction between expected inflation, real interest rates and demand conditions. Barsky et al. (2014) emphasise that the natural rate can guide policy analysis only when its measurement limits are acknowledged. In the present paper, the natural-rate gap is used as a stance indicator, but the empirical interpretation remains conditional on the uncertainty surrounding both the natural rate and the expectations-based real-rate input.

1.3. Structural transmission in a small open economy

The second strand of literature studies monetary-policy transmission with vector autoregressive and structural vector autoregressive models. Sims (1980) established VARs as a flexible empirical framework for dynamic macroeconomic analysis, while Christiano et al. (1999) show why impulse responses to monetary-policy innovations remain informative for evaluating transmission mechanisms. Kilian and Lütkepohl (2017) provide the standard framework for identifying structural shocks and interpreting impulse responses and forecast error variance decompositions. The SVEC approach extends this logic to cointegrated systems in which long-run relations and short-run dynamics are modelled jointly.

Small open economies add additional complications. External interest rates, exchange rates, imported price pressures and sovereign or country-risk premia can affect domestic transmission. Stefański (2023) studies natural-rate estimation for Central and Eastern European economies and highlights the importance of open-economy channels. Segura-Rodriguez (2026) provides related evidence for a small open economy and shows that neutral-rate estimates are sensitive to external conditions.

For Poland, benchmark estimates provide an essential comparison because country-specific models may differ in level, sample endpoint and treatment of post-pandemic shocks. The present paper uses this evidence as an external benchmark for the level and time path of the estimated Polish natural rate.

2. Research method and data

2.1. Data, transformations and empirical design

The empirical sample covers 2010Q1–2025Q1. The baseline state-space information set uses seasonally adjusted industrial production, core inflation, a short-term policy-rate proxy and survey-based inflation expectations. Robustness variants replace industrial production with quarterly GDP and GVA indicators. The SVEC block is specified in scale-consistent macro-financial transformations: log growth rates, inflation rates, spreads, percentage-point gaps and ratios.

The starting point is determined by the real-time construction of the ex ante real policy rate. The Carlson-Parkin survey indicator is converted into inflation units through a recursive scaling equation in which the coefficients are estimated from information available before the quarter being evaluated. The observations

preceding 2010Q1 provide the initial history for this recursive mapping. Beginning the state-space estimation in 2010Q1 ensures that the ex ante real-rate input enters the measurement system after the survey-based expectations series has been initialised under the same real-time convention used throughout the sample. The 2010Q1–2025Q1 window is then used consistently for the estimation of r^* , the construction of NRG and the subsequent SVEC transmission analysis.

The empirical design follows a two-stage structure. The first stage estimates the natural real interest rate and constructs the natural-rate gap. The second stage treats the natural-rate gap as an endogenous transmission variable in an SVEC framework with domestic and external weakly exogenous controls. This separation keeps the measurement problem and the transmission problem empirically distinct.

2.2. Ex ante real policy rate and inflation expectations

Let the short-term nominal policy-rate proxy be denoted by i_t and let $\hat{\pi}_{t+4|t}^{e,CP}$ denote the recursively scaled Carlson-Parkin estimate of four-quarter-ahead inflation expected in quarter t . The ex ante real policy rate is defined as:

$$r_t^{\text{ex ante}} = i_t - \hat{\pi}_{t+4|t}^{e,CP} \quad (1)$$

where:

- $r_t^{\text{ex ante}}$ – the ex ante real policy rate in quarter t ,
- i_t – the short-term nominal policy-rate proxy in quarter t ,
- $\hat{\pi}_{t+4|t}^{e,CP}$ – the recursively scaled Carlson-Parkin estimate of four-quarter-ahead inflation expectations formed in quarter t .

Let p_F, p_S, p_L, p_U, p_D and $p_{DK,t}$ denote the survey shares assigned to faster price growth, the same rate of price growth, slower price growth, unchanged prices, falling prices and non-informative answers, respectively. After excluding non-informative answers, the informative shares are normalised as:

$$\tilde{p}_{j,t} = 100 \frac{p_{j,t}}{100 - p_{DK,t}}, \quad j \in \{F, S, L, U, D\} \quad (2)$$

where:

- j – the informative response category,
- $\tilde{p}_{j,t}$ – the renormalised share of category j in quarter t ,
- $p_{DK,t}$ – the share of non-informative answers in quarter t .

The baseline Carlson-Parkin indicator is based on the renormalised directional balance:

$$B_t^{\text{CP}} = \tilde{p}_{F,t} + \tilde{p}_{S,t} + \tilde{p}_{L,t} - \tilde{p}_{D,t} \quad (3)$$

where:

B_t^{CP} – the Carlson-Parkin directional balance in quarter t .

The balance is transformed into a standardised survey indicator:

$$z_t^{\text{CP}} = \Phi^{-1} \left[\min \left(\max \left(\frac{B_t^{\text{CP}} / 100 + 1}{2}, 10^{-5} \right), 1 - 10^{-5} \right) \right] \quad (4)$$

where:

z_t^{CP} – the standardised Carlson-Parkin survey indicator in quarter t ,
 $\Phi^{-1}(\cdot)$ – the inverse cumulative distribution function of the standard normal distribution,
 10^{-5} and $1 - 10^{-5}$ – numerical bounds keeping the transformed probability inside the open unit interval.

The recursive scaling equation maps this standardised survey indicator into the inflation units used in the ex ante real-rate construction. For quarter t , the scaling regression is estimated on the historical information set T_t^{hist} :

$$\pi_s^c = a_{0,t}^{\text{CP}} + b_{1,t}^{\text{CP}} z_s^{\text{CP}} + u_s^{\text{CP}}, \quad s = 1, \dots, t-1 \quad (5)$$

where:

s – the index of historical quarters used to estimate the recursive scaling equation for quarter t ,
 T_t^{hist} – the historical information set available before quarter t ,
 π_s^c – core inflation in historical quarter s ,
 $a_{0,t}^{\text{CP}}$ – the intercept of the recursive Carlson-Parkin scaling regression for quarter t ,
 $b_{1,t}^{\text{CP}}$ – the slope coefficient of the recursive Carlson-Parkin scaling regression for quarter t ,
 z_s^{CP} – the Carlson-Parkin survey indicator in historical quarter s ,
 u_s^{CP} – the projection disturbance in the Carlson-Parkin scaling regression.

The recursively scaled Carlson-Parkin expectation is:

$$\hat{\pi}_{t+4t}^{e, \text{CP}} = \hat{a}_{0,t}^{\text{CP}} + \hat{b}_{1,t}^{\text{CP}} z_t^{\text{CP}} \quad (6)$$

where:

$\hat{\pi}_{t+4t}^{e, \text{CP}}$ – the Carlson-Parkin-based expectation of inflation over the next four quarters formed in quarter t ,
 $\hat{a}_{0,t}^{\text{CP}}, \hat{b}_{1,t}^{\text{CP}}$ – estimates of the recursive Carlson-Parkin scaling coefficients obtained from T_t^{hist} ,
 z_t^{CP} – the Carlson-Parkin survey indicator observed in quarter t .

This convention makes the expectations input real-time: the expectation formed in quarter t is scaled with information available before quarter t . When the historical window is too short or the survey indicator has negligible variation, the recursive historical mean of core inflation provides the initial scaled value.

The ordered-survey robustness measure uses the full ordinal response structure:

$$S_t^{\text{ord}} = 2\tilde{p}_{F,t} + \tilde{p}_{S,t} + 0,5\tilde{p}_{L,t} - \tilde{p}_{D,t}, \quad z_t^{\text{ord}} = \frac{S_t^{\text{ord}} - \text{mean}(S^{\text{ord}})}{\text{sd}(S^{\text{ord}})} \quad (7)$$

where:

- s_t^{ord} – the ordered-survey score in quarter t ,
- z_t^{ord} – the standardised ordered-survey score,
- $\text{mean}(S^{\text{ord}})$ – the sample mean of the ordered-survey score,
- $\text{sd}(S^{\text{ord}})$ – the sample standard deviation of the ordered-survey score.

The corresponding recursively scaled expectation is:

$$\hat{\pi}_{t+4t}^{e, \text{ord}} = \hat{a}_{0,t}^{\text{ord}} + \hat{b}_{1,t}^{\text{ord}} z_t^{\text{ord}} \quad (8)$$

where:

- $\hat{\pi}_{t+4t}^{e, \text{ord}}$ – the ordered-survey-based expectation of inflation over the next four quarters formed in quarter t ,
- $\hat{a}_{0,t}^{\text{ord}}, \hat{b}_{1,t}^{\text{ord}}$ – recursively estimated intercept and slope coefficient for the ordered-survey index.

The model-implied expectations benchmark is generated within the state-space measurement system by linking the nominal policy-rate proxy to the latent trend-inflation component and a latent real-rate process:

$$i_t = \pi_t^* + r_t^{\text{model}} + \varepsilon_{i,t} \quad (9)$$

where:

- i_t – the nominal policy-rate proxy,
- π_t^* – latent systematic inflation,
- r_t^{model} – the model-implied latent real-rate component,
- $\varepsilon_{i,t}$ – the measurement disturbance in the nominal-rate equation.

The resulting model-implied real-rate component generates an alternative NRG path used as an expectations-channel diagnostic benchmark. These alternatives are retained as robustness checks for the expectations channel, which is evaluated in Section 3.2 and documented in Figure A2.

2.3. Natural-rate state-space model

The natural-rate gap is defined as the difference between the ex ante real policy rate and the latent natural real interest rate:

$$\text{NRG}_t = r_t^{\text{ex ante}} - r_t^* \quad (10)$$

where:

NRG_t – the natural-rate gap in quarter t ,

$r_t^{\text{ex ante}}$ – the ex ante real policy rate,

r_t^* – the latent natural real interest rate estimated by the state-space model.

A positive NRG_t indicates a restrictive stance in real-rate terms, while a negative NRG_t indicates an accommodative stance.

The natural real interest rate follows the measurement decomposition:

$$r_t^* = g_t + z_t^r \quad (11)$$

where:

g_t – the trend-growth anchor of the natural-rate process,

z_t^r – the residual low-frequency component of the natural real interest rate.

The observed measurement vector is:

$$Y_t = (y_t, \pi_{c,t}, r_{m,t}, \Delta y_t)' \quad (12)$$

where:

Y_t – the vector of observed variables entering the state-space measurement system,

y_t – centred log activity,

$\pi_{c,t}$ – core inflation,

$r_{m,t}$ – the measured ex ante real-rate input,

Δy_t – quarterly activity growth.

The latent state vector is:

$$s_t = (x_t, \tau_t, g_t, r_t^*, \pi_t^*, r_{\text{exo},t})' \quad (13)$$

where:

s_t – the latent state vector,

x_t – the activity gap,

τ_t – trend activity,

g_t – annualised trend growth,

r_t^* – the natural real interest rate,

π_t^* – systematic inflation,

$r_{\text{exo},t}$ – the latent real-rate anchor corresponding to the measured ex ante real-rate input.

The measurement equations are:

$$y_t = x_t + \tau_t + \varepsilon_{y,t} \quad (14)$$

$$\pi_{c,t} = \pi_t^* + \varepsilon_{\pi,t} \quad (15)$$

$$r_{m,t} = r_{\text{exo},t} + \varepsilon_{r,t} \quad (16)$$

$$\Delta y_t = \frac{g_t}{4} + \varepsilon_{\Delta y,t} \quad (17)$$

where:

$\varepsilon_{y,t}, \varepsilon_{\pi,t}, \varepsilon_{r,t}, \varepsilon_{\Delta y,t}$ – measurement disturbances,
 $g_t / 4$ – annualised trend growth converted into a quarterly growth rate.

The measurement-error covariance matrix is:

$$H = \text{diag}(\sigma_y^2, \sigma_\pi^2, \sigma_r^2, \sigma_{\Delta y}^2) \quad (18)$$

where:

H – the covariance matrix of measurement disturbances,
 $\text{diag}(\cdot)$ – a diagonal matrix,
 $\sigma_y^2, \sigma_\pi^2, \sigma_r^2, \sigma_{\Delta y}^2$ – the variances of the measurement disturbances in the activity, inflation, real-rate and activity-growth measurement equations.

The transition equations are:

$$x_t = \phi_y x_{t-1} - a_{\text{IS}}(r_{\text{exo},t-1} - r_{t-1}^*) + \eta_{x,t} \quad (19)$$

$$\tau_t = \tau_{t-1} + \frac{g_t}{4} + \eta_{\tau,t} \quad (20)$$

$$g_t = g_{t-1} + \eta_{g,t} \quad (21)$$

$$r_t^* = r_{t-1}^* + \eta_{r^*,t} \quad (22)$$

$$\pi_t^* = b_{\text{PC}} x_{t-1} + \phi_\pi \pi_{t-1}^* + \eta_{\pi^*,t} \quad (23)$$

$$r_{\text{exo},t} = r_{\text{exo},t-1} + \eta_{r,t} \quad (24)$$

where:

ϕ_y – the persistence parameter of the activity gap,
 a_{IS} – the semi-elasticity of the activity gap with respect to the lagged real-rate gap,
 $\eta_{x,t}$ – the activity-gap state innovation,
 $\eta_{\tau,t}$ – the trend-activity innovation,
 $\eta_{g,t}$ – the trend-growth innovation,
 $\eta_{r^*,t}$ – the natural-rate innovation,

- b_{PC} – the Phillips-curve slope with respect to the lagged activity gap,
 ϕ_π – the persistence parameter of systematic inflation,
 $\eta_{\pi^*,t}$ – the systematic-inflation innovation,
 $\eta_{r,t}$ – the innovation to the latent real-rate anchor.

The state-innovation covariance matrix is:

$$Q = \text{diag}(\sigma_x^2, \sigma_\tau^2, \sigma_g^2, \sigma_{r^*}^2, \sigma_{\pi^*}^2, \sigma_r^2) \quad (25)$$

where:

- Q – the covariance matrix of state innovations,
 $\sigma_x^2, \sigma_\tau^2, \sigma_g^2, \sigma_{r^*}^2, \sigma_{\pi^*}^2, \sigma_r^2$ – the innovation variances of the activity gap, trend activity, trend growth, natural real interest rate, systematic inflation and latent real-rate anchor.

The measured ex ante real-rate state anchors the observed real-rate input in the measurement block and enters the IS-type propagation with a one-quarter lag. This timing convention treats the real-rate gap as predetermined for the current activity-gap transition.

The model is estimated by Gaussian maximum likelihood using the Kalman filter. The baseline path is obtained from the two-sided Kalman smoother:

$$E(s_t | \mathcal{F}_T) \quad (26)$$

where:

- $E(s_t | \mathcal{F}_T)$ – the conditional expectation of the latent state vector in quarter t using the full-sample information set $\mathcal{F}_T$.

The end-of-sample robustness exercise uses the one-sided filtered state estimate:

$$E(s_t | \mathcal{F}_t) \quad (27)$$

where:

- $E(s_t | \mathcal{F}_t)$ – the conditional expectation of the latent state vector in quarter t using information available up to quarter t .

The filtered estimate is extracted from the same Kalman recursion and compared with the smoothed estimate over the common sample. The system combines an IS-type activity-gap equation, a Phillips-curve-type inflation equation, low-frequency trend processes and a measured real-rate anchor. Smoothed state estimates are used to recover r_t^* and the natural-rate gap, while filtered estimates are retained for the endpoint-sensitivity exercise.

2.4. Robustness design for the natural real interest rate

The robustness design has four components. First, the baseline survey-expectations series is compared with the ordered-survey index and the model-implied benchmark described above. Second, the activity measure is varied by replacing industrial production with GDP and GVA, which checks whether the estimated natural-rate path is driven by the narrower industrial proxy. Third, selected Q and H variances are perturbed: tight, loose and smooth variants test moderate signal-to-noise sensitivity, while the flexible- r^* variant deliberately relaxes the low-frequency natural-rate process and is treated as a stress test. Fourth, the baseline smoother is compared with a one-sided filter to assess how much of the endpoint profile depends on two-sided information.

The robustness exercises are interpreted asymmetrically. Variants that preserve the baseline information set but alter a measurement choice are treated as robustness checks. Highly flexible natural-rate decompositions are treated as diagnostic stress tests if they generate implausible low-frequency paths. This distinction is important because the empirical objective is not to maximise variation in r^* , but to obtain a disciplined stance measure suitable for subsequent transmission analysis.

2.5. SVEC transmission framework

Let x_t denote the vector of endogenous transmission variables and q_t the vector of weakly exogenous controls. The baseline endogenous transmission vector is:

$$x_t = (\text{activity}_t, \text{CPI}_t, \text{REER}_t, \text{term_spread}_t, \text{debt_serv_cost}_t, \text{NRG}_t)' \quad (28)$$

where:

- activity_t – seasonally adjusted industrial-production growth,
- CPI_t – consumer price inflation,
- REER_t – the real effective exchange rate,
- term_spread_t – the term spread between the long-term and short-term government-bond yield,
- debt_serv_cost_t – debt-service costs measured as a four-quarter ratio,
- NRG_t – the natural-rate gap.

The extended robustness specification augments the system with the unemployment rate. The activity-measure robustness scenarios replace industrial-production growth with GDP growth and GVA growth, respectively, while keeping the remaining SVEC variables and weakly exogenous controls unchanged. The weakly exogenous control vector is:

$$q_t = (m_t, \text{banks_profit}_t, \text{EA_gdp}_t, \text{ECB_deposit}_t, \text{Brent}_t, \text{D_covid}_t, \text{D_energy}_t)' \quad (29)$$

where:

- q_t – the vector of weakly exogenous controls,
- m_t – the M1-to-GDP ratio,
- banks_profit_t – banking-sector net profits relative to interest costs,
- EA_gdp_t – euro-area real GDP growth,
- ECB_deposit_t – the ECB deposit facility rate,
- Brent_t – Brent oil-price growth,
- D_covid_t – the pandemic-period dummy,
- D_energy_t – the energy-war-shock dummy.

The controls enter outside the cointegration space and absorb observable domestic financial conditions, external demand, foreign monetary-policy conditions, global energy-price pressure and intervention episodes. The reduced-form VAR in levels can be written as:

$$x_t = c + A_1 x_{t-1} + \dots + A_p x_{t-p} + \Psi q_t + u_t \quad (30)$$

where:

- c – the deterministic component,
- $A_1, \dots, A_p$ – autoregressive coefficient matrices,
- Ψ – the coefficient matrix on weakly exogenous controls,
- u_t – the reduced-form innovation vector,
- p – the VAR lag order.

When the endogenous variables are cointegrated, the system admits the VECM representation:

$$\Delta x_t = \alpha \beta' x_{t-1} + \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{p-1} \Delta x_{t-p+1} + \Psi q_t + u_t \quad (31)$$

where:

- Δx_t – the first difference of the endogenous transmission vector,
- α – the loading matrix,
- β – the matrix of cointegrating vectors,
- $\Gamma_1, \dots, \Gamma_{p-1}$ – short-run coefficient matrices,
- p – the VAR lag order carried into the VECM representation.

The integration order of the SVEC variables is assessed using ADF and Phillips-Perron tests with the unit-root null and KPSS tests with the stationarity null. The baseline activity variable is industrial-production growth because the log level of industrial production displays an integration profile incompatible with the baseline cointegrated system. Lag length is evaluated with standard VAR information criteria for dynamic systems as discussed by Lütkepohl (2005), with the lag-order diagnostics across SVEC scenarios reported in Figure C2. The cointegration rank is

selected using the likelihood-based trace-testing framework of Johansen (1995). The selected lag and cointegration-rank structure is reported in Appendix C.

Structural shocks are recovered from reduced-form innovations with zero mean and covariance matrix Σ_u . The structural representation is defined as:

$$u_t = B\varepsilon_t, \quad E(\varepsilon_t \varepsilon_t') = I \quad (32)$$

where:

- B – the contemporaneous structural impact matrix,
- ε_t – the vector of orthonormal structural shocks,
- I – the identity matrix,
- $E(\cdot)$ – the expectation operator,
- Σ_u – the covariance matrix of reduced-form innovations.

The baseline identification uses a recursive short-run structure. The natural-rate-gap shock is scaled so that its impact corresponds to a 25-basis-point increase in the NRG. The same scaling is used for impulse responses, accumulated responses and FEVD summaries, which are evaluated in Section 3.3. Impulse responses are interpreted as responses to a model-conditional stance innovation, not as a complete structural monetary-policy shock.

3. Results and discussion

3.1. Baseline natural real interest rate and monetary-policy stance

Figure 1 shows the baseline estimate of Poland's natural real interest rate together with the 95% state-uncertainty interval. The estimated mean of r^* over the full sample is 0.87 percentage points, with a standard deviation of 1.34 percentage points, a minimum of -1.70 and a maximum of 3.55 . The profile combines a gradual decline through the low-inflation period, a pandemic trough and a sharp post-pandemic increase. The average confidence-interval width of 0.77 percentage points is large enough to make interval estimates more informative than point estimates, especially around the pandemic and tightening episodes.

Figure A1 reports the observed series entering the state-space measurement block of the LW model and the ex ante real policy-rate construction. The estimated level of r^* is broadly consistent with Polish benchmark evidence, including the range and post-pandemic direction reported by Bielecki et al. (2023). The post-2021 increase is economically coherent with the tightening of the ex ante real policy rate and the recovery of activity after the pandemic shock.

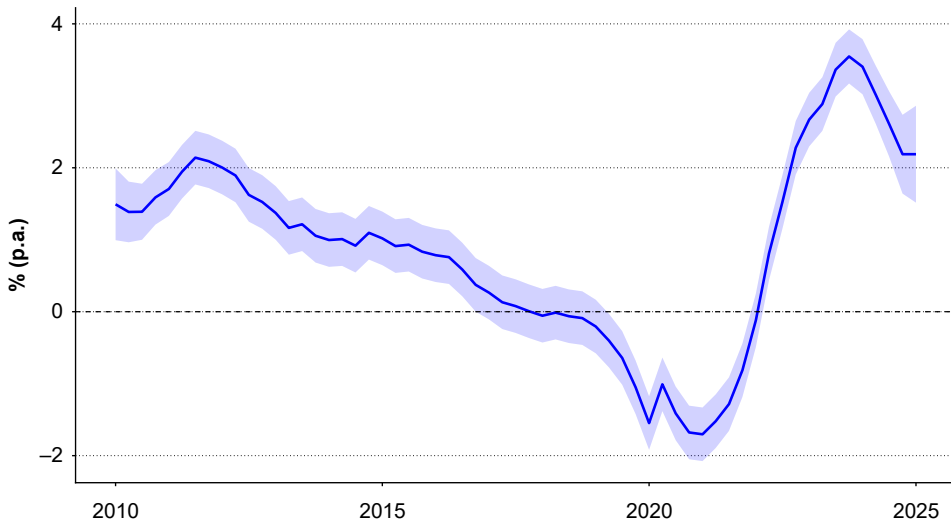


Figure 1. Natural real interest rate (r^*) with 95% confidence interval

Source: author's analysis.

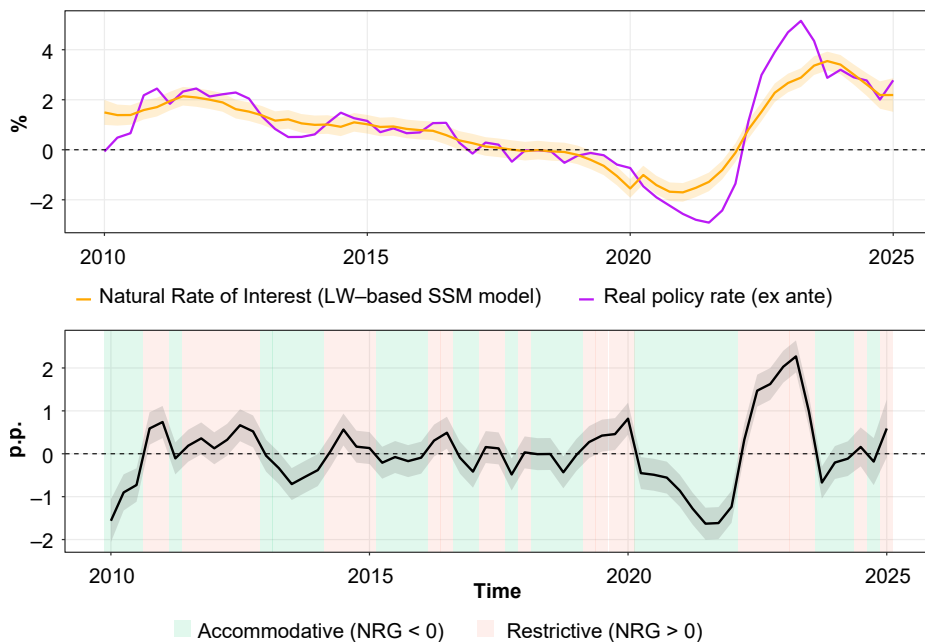


Figure 2. Monetary policy stance: NRG and ex ante real policy rate

Source: author's analysis.

Figure 2 translates the natural-rate estimate into a monetary-policy stance indicator. The natural-rate gap has an average close to zero over the full sample, while restrictive observations account for 47.5% of quarters in the baseline specification. This near balance hides two very different subperiods: the pandemic and immediate post-pandemic quarters are dominated by negative real-rate gaps, and the tightening cycle generates the strongest positive gaps in the sample. The figure identifies the timing of stance changes more clearly than a nominal-rate plot would.

Table 1 provides the numerical counterpart to Figures 1 and 2. The baseline and ordered-survey variants are close: their mean r^* estimates are 0.87 and 0.84 percentage points, and the ordered-survey path remains highly correlated with the baseline. The model-implied expectations benchmark is the main outlier, with a mean r^* of -0.91 percentage points, a larger standard deviation of 1.84 and a restrictive-stance share of 75.4%. Survey aggregation leaves the central stance narrative stable; replacing survey expectations with a model-implied proxy introduces substantially greater measurement sensitivity.

Table 1. Estimated r^* and NRG across selected specifications

Specification	Mean r^*	Min r^*	Max r^*	Mean NRG	Min NRG	Max NRG	Restrictive quarters
Baseline	0.87	-1.70	3.55	0.01	-1.63	2.27	47.5%
Model-implied expectations	-0.91	-4.91	1.37	0.86	-1.94	4.29	75.4%
Ordered-survey expectations	0.84	-1.43	3.74	0.01	-1.59	2.37	44.3%
GDP activity	0.91	-1.61	3.76	-0.03	-1.69	1.98	49.2%
GVA activity	0.90	-1.63	3.73	-0.02	-1.68	1.96	49.2%

Note: r^* and NRG are expressed in percentage points. NRG is defined as the ex ante real policy rate minus r^* . Restrictive quarters denote observations with $\text{NRG} > 0$.

Source: author's calculations.

3.2. Robustness of r^* and the natural-rate gap

Figure 3 evaluates the expectations channel. The ordered-survey variant tracks the baseline closely: the correlation with the baseline r^* path is 0.975, the mean absolute deviation is 0.25 percentage points and the maximum deviation is 0.55 percentage points. Figure A2 explains this result visually: the Carlson-Parkin and ordered-survey measures contain similar low-frequency information because

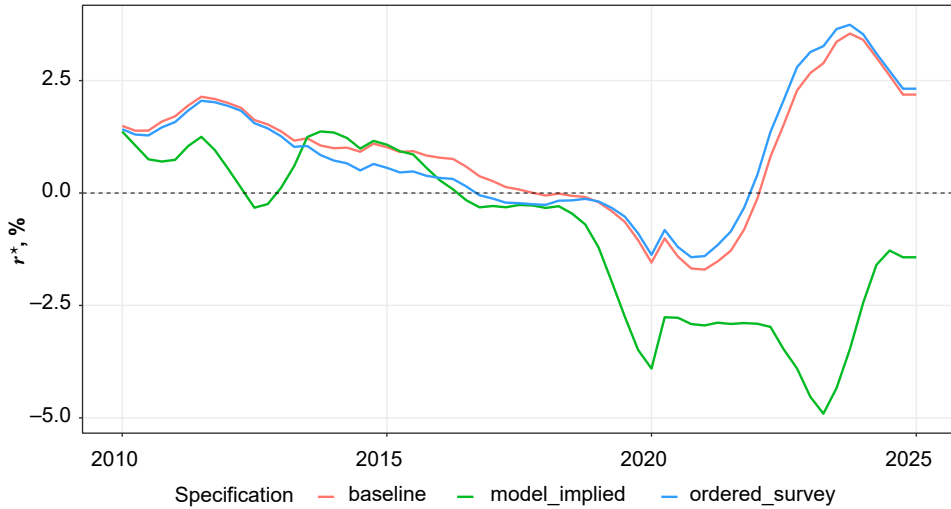


Figure 3. Robustness of r^* to alternative inflation-expectations measures

Source: author's analysis.

both are anchored in the same qualitative survey, but the ordered-survey index preserves more information from the response distribution. The model-implied expectations benchmark behaves differently, with a correlation of only 0.148 and a mean absolute r^* deviation of 1.82 percentage points.

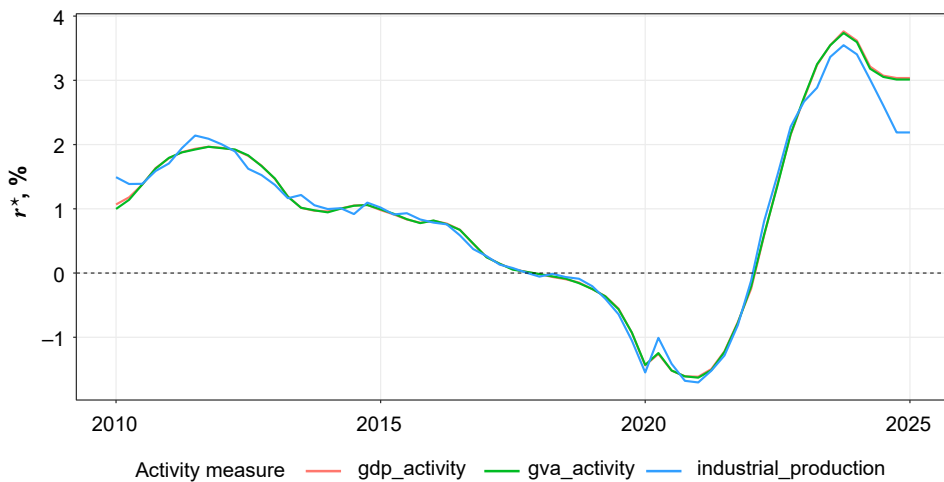


Figure 4. Robustness of r^* to alternative activity measures

Source: author's analysis.

Figure 4 addresses the activity-measure concern. Replacing industrial production with GDP or GVA leaves the broad r^* profile intact: correlations with the baseline are 0.99 for both activity alternatives, and mean absolute differences are only about 0.13 percentage points. The GDP and GVA variants also remain inside the baseline 95% state-uncertainty interval in 96.7% of common observations. Industrial production affects short-run timing, and the GDP/GVA profiles support the use of the baseline activity proxy.

The ordered-survey and model-implied variants modify the expectations block. The GDP and GVA variants modify the activity block. The tight, loose and smooth specifications modify the state-space calibration, while the flexible- r^* variant is treated as a diagnostic stress test. This classification matters because these variants do not have the same evidential status. A specification that changes the activity proxy is a robustness check of the baseline information set; a specification that allows excessive low-frequency movement in r^* is used to test the limits of identification. The calibration results shown in Figure 5 support this distinction. Tight and loose growth-anchor specifications differ from the baseline by only 0.005 and 0.015 percentage points on average, and both remain inside the baseline confidence interval throughout the sample. The smooth- r^* variant is also close, with

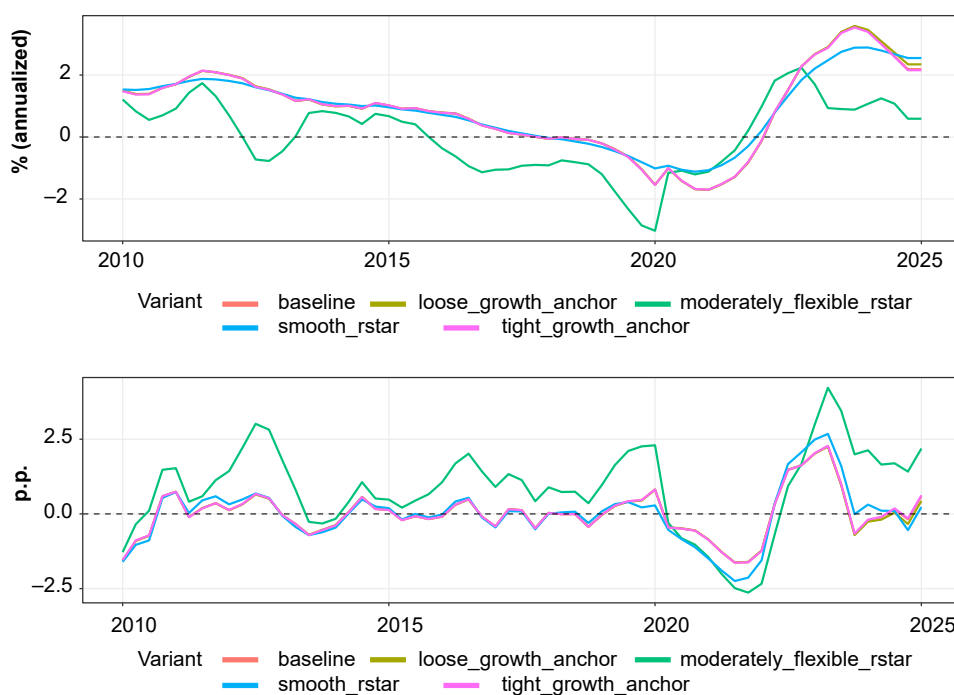


Figure 5. Calibration-variance robustness of r^* and NRG

Source: author's analysis.

a mean absolute difference of 0.19 percentage points and a correlation of 0.987. In contrast, the moderately flexible r^* specification has a mean absolute difference of 1.05 percentage points and remains inside the baseline interval in only 14.8% of observations. The stress version is much more extreme, with a mean absolute difference of 8.06 percentage points. These diagnostics justify using the flexible variants as stress tests, not as alternative central estimates.

The calibration-variance evidence in Figure 5 shows a stable central cluster and one diagnostic outlier class. Tight, loose and smooth calibration variants preserve the timing of the pandemic trough and the post-pandemic rise, indicating that the main r^* profile is not driven by a narrow choice of signal-to-noise ratios. Highly flexible decompositions absorb excessive low-frequency movement into r^* and produce paths that are difficult to reconcile with the survey-based real-rate input. The empirical interpretation is based on the baseline, ordered-survey, activity and moderate-calibration specifications; the highly flexible variants are reported as stress tests.

Figure 6 compares the two-sided smoother with the one-sided filter. The two estimates share the same broad cycle, but the filtered path is more jagged at the endpoint because it uses information available only up to each quarter. This distinction is central for real-time policy interpretation: it shows that the post-2021 increase is visible under both extraction methods, while the exact level of the latest observations is more sensitive to the filtering convention. Figure B1 shows that the endpoint sensitivity coincides with large movements in the latent output gap and trend-inflation components, and Figure 5 confirms that moderate cali-

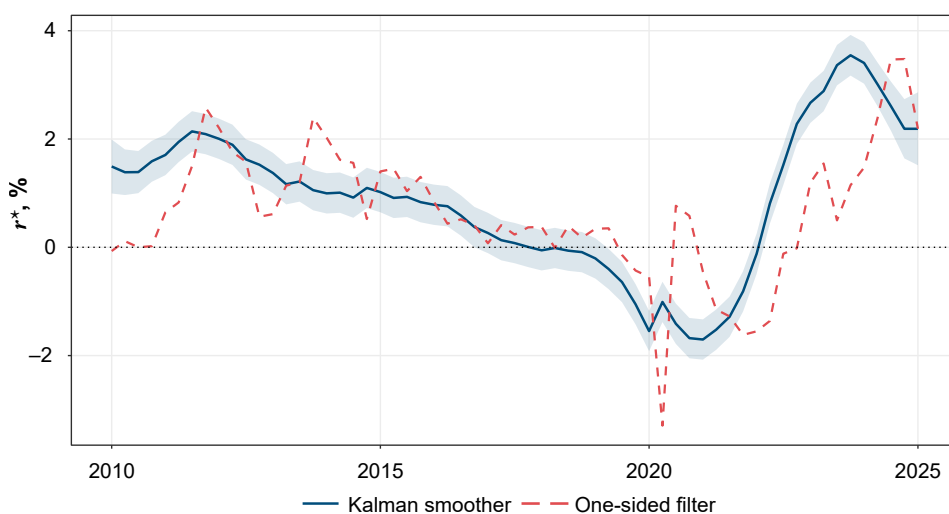


Figure 6. End-of-sample behaviour of r^* : one-sided filter versus smoother

Source: author's analysis.

bration changes do not overturn the baseline profile. The post-pandemic rise in r^* is therefore not an artefact of one graphical smoothing convention, although its exact endpoint remains uncertain.

Table B1 summarises the baseline and alternative state-space specifications. Its purpose is to separate robustness exercises that preserve the economic content of the baseline from diagnostic stress tests that deliberately loosen the natural-rate process. The evidence is strongest where alternative specifications preserve the same economic content as the baseline: ordered-survey expectations, GDP/GVA activity and moderate calibration variants. The evidence is weaker where the specification changes the role of expectations or gives r^* excessive low-frequency flexibility. The measurement conclusion is therefore precise: the direction and timing of the baseline r^* cycle are robust, while the level of r^* is most sensitive to the expectations block and to the variance assigned to the natural-rate component.

3.3. SVEC transmission evidence

The SVEC stage evaluates whether innovations to the natural-rate gap contain systematic information for macro-financial transmission after the measurement uncertainty surrounding r^* has been explicitly carried into the empirical design. The baseline specification uses a parsimonious six-variable transmission block consisting of activity growth, CPI inflation, the real effective exchange rate, the term spread, debt-service costs and the natural-rate gap. The model is estimated on 58 quarterly observations and includes seven weakly exogenous domestic and external controls. Figure C1 documents the underlying transmission variables and alternative NRG measures used across the SVEC robustness scenarios.

Table C1 shows that the SVEC transmission block is empirically persistent: activity growth and the natural-rate gap are classified as stationary, while inflation, the real effective exchange rate, the term spread, unemployment and the ex ante real policy rate are treated as $I(1)$ or weakly $I(1)$, and the log level of industrial production and debt-service costs remain ambiguous across tests. This mixed diagnostic profile supports the specification in which activity growth is used instead of the log activity level, while the cointegration-based framework is retained for the persistent macro-financial block. The rank diagnostics are stable across the main robustness scenarios: with 58 observations, seven controls and two VAR lags, the selected rank is three in the baseline and in almost all alternative NRG/activity specifications, with only the extended baseline selecting rank four.

The baseline model uses the BIC-selected lag order for the reduced-form dynamics, with the Johansen step evaluated at lag order $p = 2$ and three selected cointegrating relations. The NRG lower-bound, NRG upper-bound, model-implied

expectations, GDP-activity and GVA-activity scenarios preserve the same rank, which supports comparability across the main robustness exercises.

The loading coefficients indicate adjustment mainly through activity, CPI inflation, REER, the term spread and the NRG equation, while the debt-service-cost equation displays very small adjustment coefficients. In the baseline VECM, the activity equation adjusts strongly to the first and third cointegrating relations, CPI inflation adjusts to the first relation, REER and the term spread adjust mainly to the third relation, and the NRG equation adjusts to the first two relations. These coefficients describe error-correction behaviour in the cointegrated system and provide the long-run adjustment background for the subsequent impulse-response analysis.

The recursive short-run identification imposes 15 zero restrictions, exactly matching the requirement for the six-variable system. The NRG shock is scaled to a 25-basis-point increase in the natural-rate gap. The diagnostic profile supports a cautious interpretation. The Portmanteau statistic does not indicate residual autocorrelation in the baseline model, and the multivariate ARCH statistic is benign. The LM/BG statistic is unfavourable, and the maximum root modulus is close to unity. These features imply that the responses should be interpreted as persistent, model-conditional transmission profiles. The robustness scenarios do not remove this qualification, but they show whether the economic pattern survives alternative NRG and transmission-block constructions.

Figure 7 reports the impulse responses and accumulated responses to a 25-basis-point increase in the natural-rate gap. The activity response is non-monotonic: its largest negative response is about -0.48 at horizon 3, followed by partial rebounds, and the accumulated response at horizon 13 is close to -0.11 . CPI increases by about 0.34 at horizon 1 and remains positive across the displayed ho-

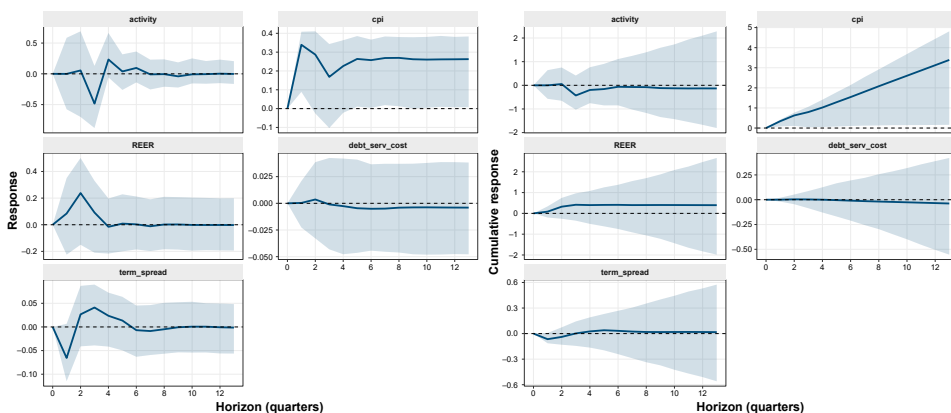


Figure 7. IRFs and accumulated IRFs to a 25-basis-point NRG shock in the baseline SVEC

Source: author's analysis.

rizon. The REER response peaks at about 0.24 at horizon 2, while term-spread and debt-service-cost responses are smaller and more oscillatory.

Table 2 condenses the IRF and FEVD evidence into sign, timing, peak and cumulative-response summaries. Statistically informative responses are sparse. CPI provides the clearest limited nominal evidence in the baseline summary, but its FEVD contribution remains modest at about 5.4–5.5%. Activity, REER, the term spread and debt-service costs display small, unstable or imprecisely estimated responses with limited FEVD shares. The baseline evidence therefore supports a restricted interpretation: the NRG innovation contains some nominal information, while it does not generate a broad and precisely estimated propagation profile across the macro-financial system.

Table 2. Summary of IRFs and FEVD for the baseline NRG shock

Response	Peak IRF	Trough IRF	Cumulative response, $h = 13$	NRG FEVD, $h = 4$	NRG FEVD, $h = 8$	NRG FEVD, $h = 13$
Activity	0.23 ($h = 4$)	−0.48 ($h = 3$)	−0.11	1.0%	1.2%	1.0%
Inflation	0.34* ($h = 1$)	0.17 ($h = 3$)	3.43*	5.4%	5.4%	5.5%
REER	0.24 ($h = 2$)	−0.02 ($h = 4$)	0.38	1.2%	0.7%	0.5%
Term spread	0.04 ($h = 3$)	−0.06 ($h = 1$)	0.02	1.9%	1.1%	0.7%
Debt-service costs	0.00 ($h = 2$)	−0.01 ($h = 6$)	−0.04	0.0%	0.0%	0.0%

Note: IRF and accumulated responses are reported for a 25-basis-point NRG shock. h denotes the horizon, measured in quarters. * denotes a response whose 95% confidence interval excludes zero.

Source: author's calculations.

Figure 8 reports the forecast-error-variance contribution of the NRG shock. The contribution is largest for the NRG equation itself, falling from about 39.7% at horizon 4 to 17.1% at horizon 13. For CPI it is stable at about 5.4–5.5% across medium horizons, while for activity it is close to 1% and for REER, the term spread and debt-service costs it is smaller. Figure C3 gives the full FEVD across all structural shocks and shows that most forecast-error variation is explained by other innovations. The NRG shock is economically interpretable, with limited dominance in the variance decomposition.

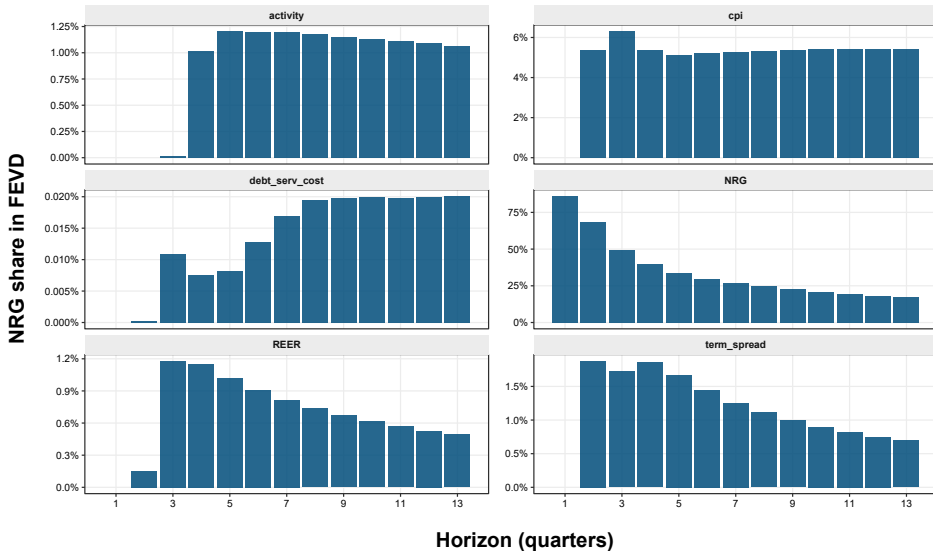


Figure 8. FEVD contribution of the NRG shock in the baseline SVEC

Source: author's analysis.

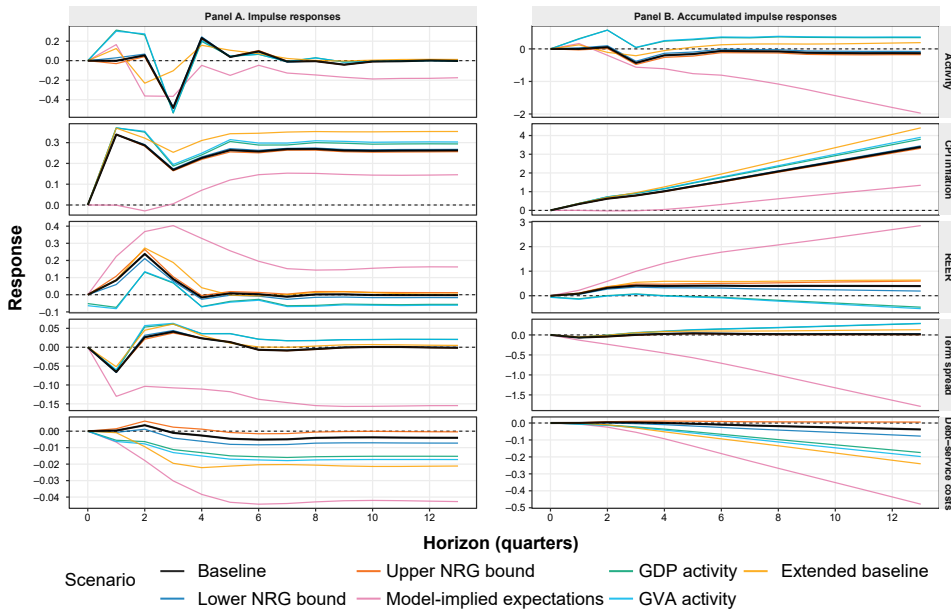


Figure 9. Robust SVEC responses to a 25-basis-point NRG shock

Source: author's analysis.

Overall, the SVEC evidence points to a bounded and statistically weak transmission role for the natural-rate gap. The identified NRG innovation can be embedded in the cointegrated macro-financial system without producing unstable dynamic profiles, and it contains limited nominal information visible mainly in CPI point estimates and modest FEVD shares. At the same time, it does not dominate the forecast-error variance of real activity, debt-service costs, REER or the term spread. The empirical contribution of the SVEC stage is therefore diagnostic and conditional: it evaluates the transmission content of an estimated stance indicator within a short quarterly sample.

Figure 9 compares the baseline SVEC responses with NRG-bound, expectations-based, activity-measure and extended-system variants. Across robustness scenarios, the sign and timing of CPI responses are more stable than the responses of other variables, although the evidence remains limited by wide confidence intervals and modest FEVD shares. The robustness exercise mainly shows that the estimated NRG innovation is not a dominant driver of activity, debt-service costs, exchange-rate dynamics or term-spread variation.

The robustness FEVD reported in Appendix Figure C4 reinforces the same ranking of effects. The NRG shock explains a limited share of forecast-error variance for activity and debt-service costs in the baseline-type specifications, a more visible but still bounded share for CPI inflation, and a strongly specification-dependent share for the term spread and REER. The own contribution of the NRG shock to NRG forecast uncertainty declines with the horizon, which indicates that the stance measure is not an isolated exogenous driver in the system but interacts with broader macro-financial dynamics over medium horizons.

4. Policy implications and limitations

The results have three policy implications. First, monetary-stance assessment for Poland should be formulated in *ex ante* real-rate terms and benchmarked against a range of natural-rate estimates, rather than inferred from nominal policy rates alone. Second, the natural-rate gap is informative as a compact measure of the intertemporal policy stance, but it should not be treated as a sufficient statistic for macroeconomic transmission. The SVEC evidence points to selective nominal and external-financial adjustment, with the clearest transmission evidence for CPI inflation and more specification-dependent responses in exchange-rate and term-spread dynamics. Third, policy communication should therefore combine the natural-rate gap with uncertainty bands, alternative expectation measures, activity-measure robustness checks and transmission diagnostics. This is particularly

important in a small open economy, where domestic monetary conditions interact with foreign demand, external interest rates, commodity-price shocks and risk premia. In this setting, the natural-rate gap is most useful as a disciplined stance indicator, not as a standalone explanation of aggregate fluctuations.

The quarterly sample is short for a structural cointegrated system and includes the pandemic, the energy-price shock and the subsequent tightening cycle. The Carlson-Parkin expectations measure is useful because it permits an ex ante real-rate construction, but it remains a survey-based transformation rather than a direct numerical expectation series. The natural real interest rate is a latent, model-dependent object, and the SVEC evidence is conditional on the estimated r^* input, the transformed macro-financial variables, the weakly exogenous controls, the selected lag/rank structure and the recursive short-run identification scheme. These limitations do not invalidate the empirical framework, but they require the results to be interpreted as conditional evidence on NRG-based transmission rather than as a complete structural account of Polish monetary-policy transmission.

Conclusions

This paper estimates Poland's time-varying natural real interest rate and natural-rate gap over 2010Q1–2025Q1 and examines the transmission of a natural-rate-gap innovation in a structural cointegrated framework. The results show that r^* is time-varying, declines around the pandemic period and rises sharply during the post-pandemic tightening episode. The implied monetary-policy stance changes materially when the ex ante real policy rate is evaluated relative to the estimated natural real interest rate.

The robustness evidence supports the central cyclical profile of r^* under alternative expectation measures, GDP/GVA activity indicators, moderate calibration variants and one-sided filtering. At the same time, the level of r^* remains sensitive to the expectations block and to highly flexible natural-rate decompositions, which are best interpreted as diagnostic stress tests rather than alternative central estimates.

The SVEC evidence indicates a bounded and statistically weak transmission role for the natural-rate gap. A 25-basis-point NRG innovation produces sparse statistically informative responses, with the clearest limited evidence in CPI inflation and small forecast-error-variance contributions for most other variables. The NRG shock does not dominate the transmission system outside the stance equation. The transmission stage therefore supports the use of the NRG as a disciplined stance indicator, while its quantitative macro-financial propagation remains conditional on the model, the short sample and the estimated r^* input.

The contribution of the paper is therefore a coherent empirical framework for measuring, testing and transmitting a natural-rate-based stance indicator in a small open economy. Future research should extend the sample, compare survey-based expectations with additional direct numerical expectations measures where available, and examine whether nominal, external-financial and fiscal-financial transmission patterns remain visible in alternative structural frameworks.

Appendix

A. Data and measurement audit

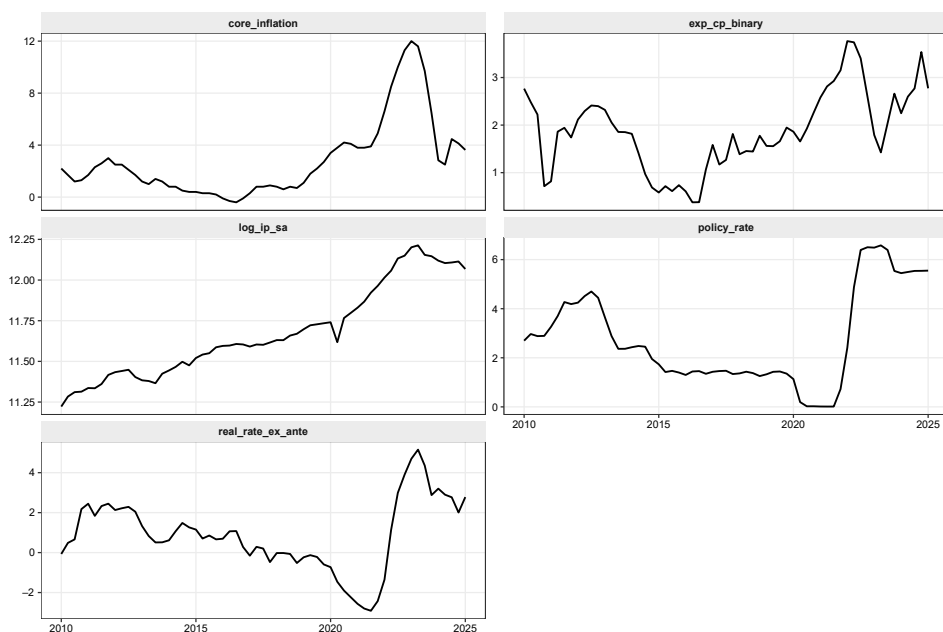


Figure A1. LW input series and ex ante real policy rate

Source: author's analysis.

Figure A2 reports the survey-based measures used to assess the robustness of inflation expectations in the LW framework. The Carlson-Parkin binary measure converts the share of respondents expecting prices to rise or fall into a quantitative expected-inflation series under the standard probability-balance interpreta-

tion. The ordered-survey score uses the full ordered structure of the survey answers, assigning higher values to stronger expected price increases and lower values to expected price decreases. The two standardised variants express these survey indicators in comparable units by removing their sample means and scaling by their sample standard deviations. These transformations allow the baseline expectation measure to be compared with a distribution-free ordered-response index and with standardised versions of both indicators.

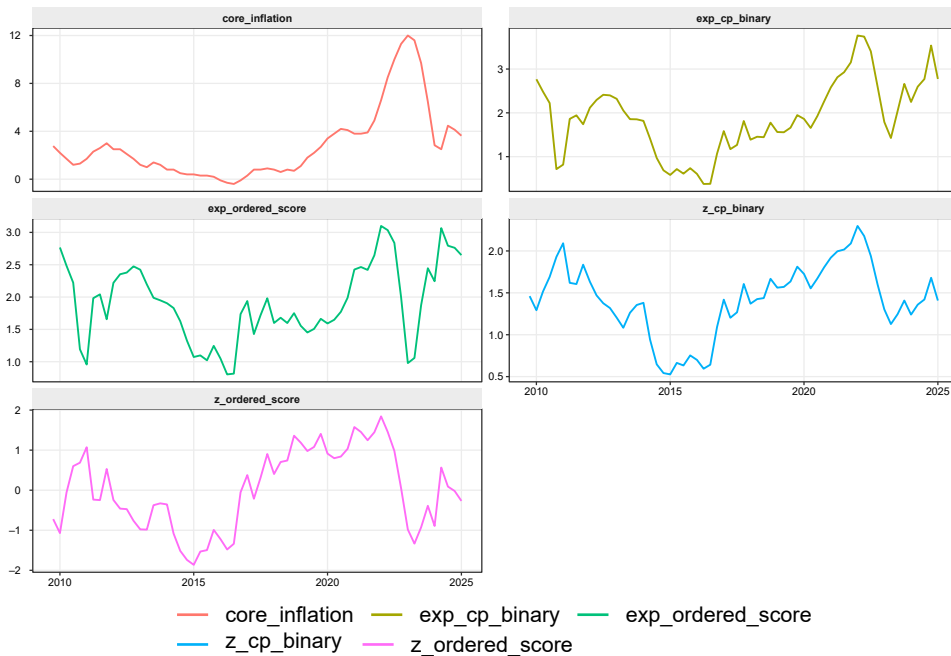


Figure A2. Carlson-Parkin expectations and ordered-survey robustness index

Source: author's analysis.

B. State-space robustness and diagnostics

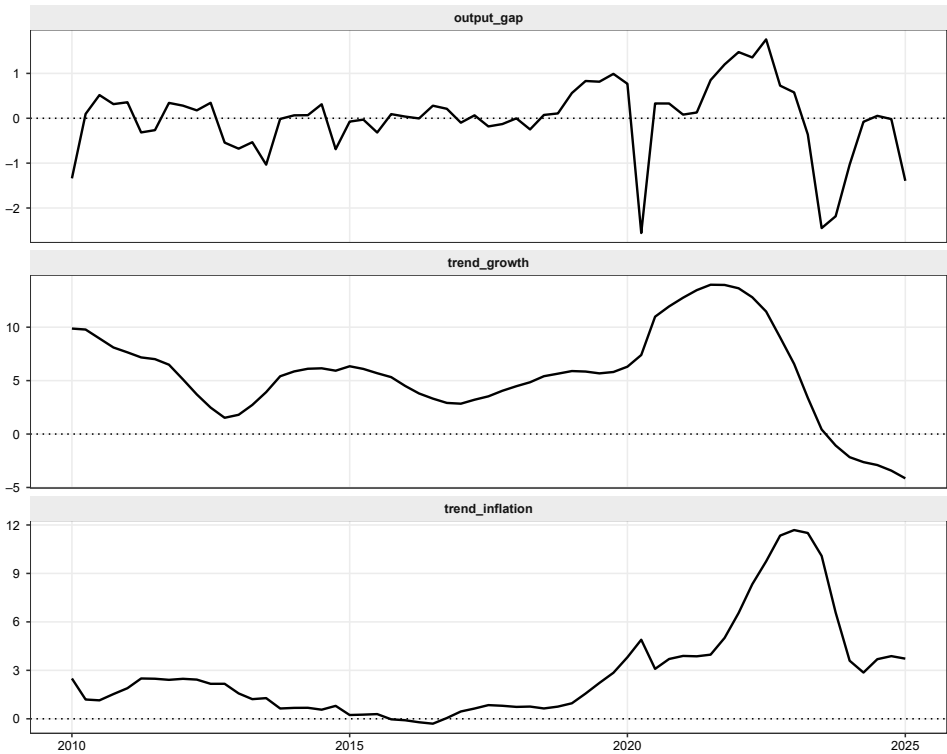


Figure B1. LW latent-state diagnostics

Source: author’s analysis.

Table B1. State-space robustness specifications

Specification	Expectation measure	Activity measure	Main modification	Empirical purpose
Baseline	Carlson-Parkin enterprise-survey index	Industrial production	Reference LW state-space model	Benchmark path for r^* and NRG
Ordered-survey expectations	Ordered enterprise-survey score	Industrial production	Full ordered response distribution	Checks whether the binary Carlson-Parkin balance drives the results
Model-implied expectations	Model-implied expected inflation	Industrial production	Alternative expectation proxy	Assesses sensitivity to expectations not taken directly from survey balances
GDP activity	Carlson-Parkin enterprise-survey index	Eurostat real GDP	Alternative activity measure	Tests robustness to a broad output aggregate
GVA activity	Carlson-Parkin enterprise-survey index	Eurostat real GVA	Alternative activity measure	Tests robustness to a production-side aggregate
Tight growth anchor	Carlson-Parkin enterprise-survey index	Industrial production	Lower trend-growth variance	Checks sensitivity to a more persistent growth component
Loose growth anchor	Carlson-Parkin enterprise-survey index	Industrial production	Higher trend-growth variance	Checks sensitivity to a more flexible growth component
Smooth r^*	Carlson-Parkin enterprise-survey index	Industrial production	Lower r^* innovation variance	Checks sensitivity to smoother natural-rate dynamics
Flexible- r^* stress test	Carlson-Parkin enterprise-survey index	Industrial production	Very high r^* innovation variance	Diagnostic stress test only

Source: author's analysis.

C. SVEC diagnostics and robustness

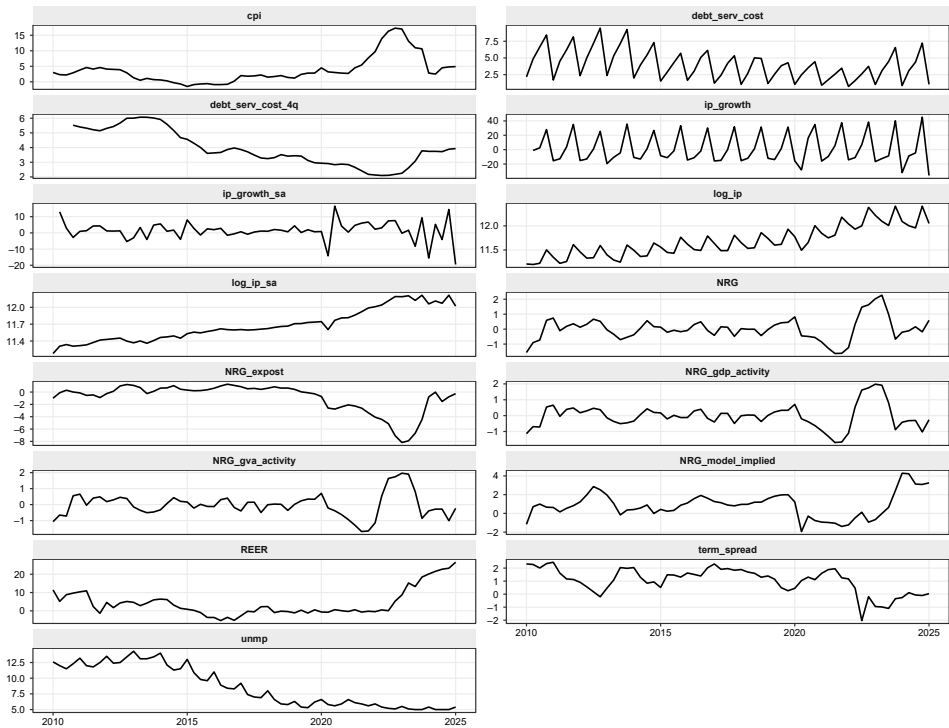


Figure C1. SVEC transmission-block variables and NRG alternatives

Source: author’s analysis.

Table C1. Unit-root, lag-order and cointegration-rank diagnostics

Panel A. Unit-root diagnostics					
Series	ADF level	PP level	KPSS level	ADF first difference	Working classification
Industrial production, log	0.907	0.010**	0.010**	0.182	Ambiguous
Industrial production growth	0.182	0.010**	0.100	0.010**	I(0)
Inflation	0.061*	0.570	0.020**	0.045**	I(1)
Real effective exchange rate	0.990	0.990	0.073*	0.010**	I(1), weak evidence
Debt-service costs	0.990	0.010**	0.010**	0.047**	Ambiguous
Term spread	0.130	0.120	0.018**	0.022**	I(1)
Unemployment	0.953	0.667	0.010**	0.636	I(1)
Ex ante real policy rate	0.356	0.664	0.100	0.083*	I(1)
Natural-rate gap	0.010**	0.026**	0.100	0.011**	I(0)
Panel B. Cointegration rank and lag-order diagnostics					
Scenario	Observations	Endogenous variables	Controls	VAR lag	Selected rank
Baseline	58	6	7	2	3
Extended baseline	58	7	7	2	4
Lower NRG bound	58	6	7	2	3
Upper NRG bound	58	6	7	2	3
Model-implied expectations	58	6	7	2	3
GDP activity	58	6	7	2	3
GVA activity	58	6	7	2	3

Note: In Panel A, ADF and PP tests use the unit-root null, whereas KPSS uses the stationarity null. ***, ** and * denote rejection at the 1%, 5% and 10% levels, respectively. Panel B reports the cointegration rank and lag order used in each SVEC scenario.

Source: author's calculations.

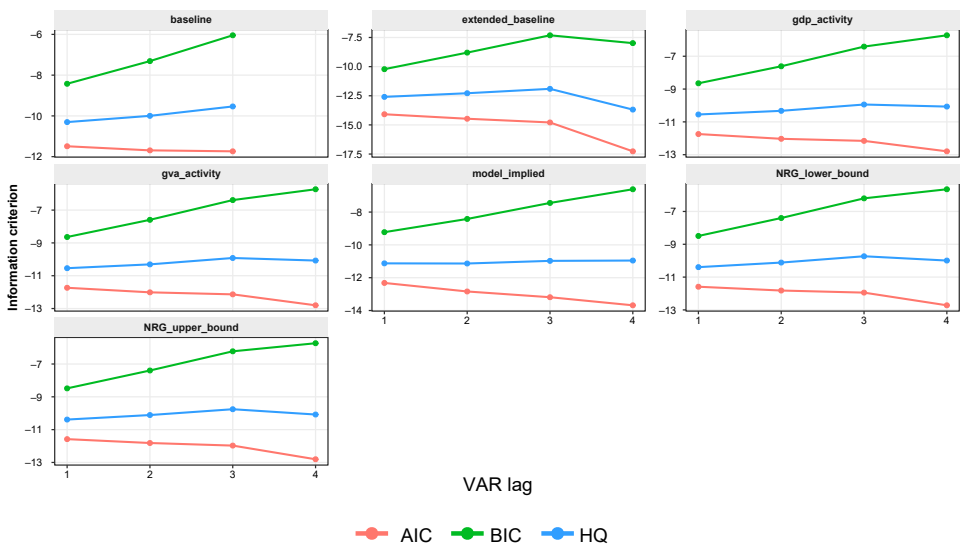


Figure C2. Lag-order selection across SVEC scenarios

Source: author’s analysis.

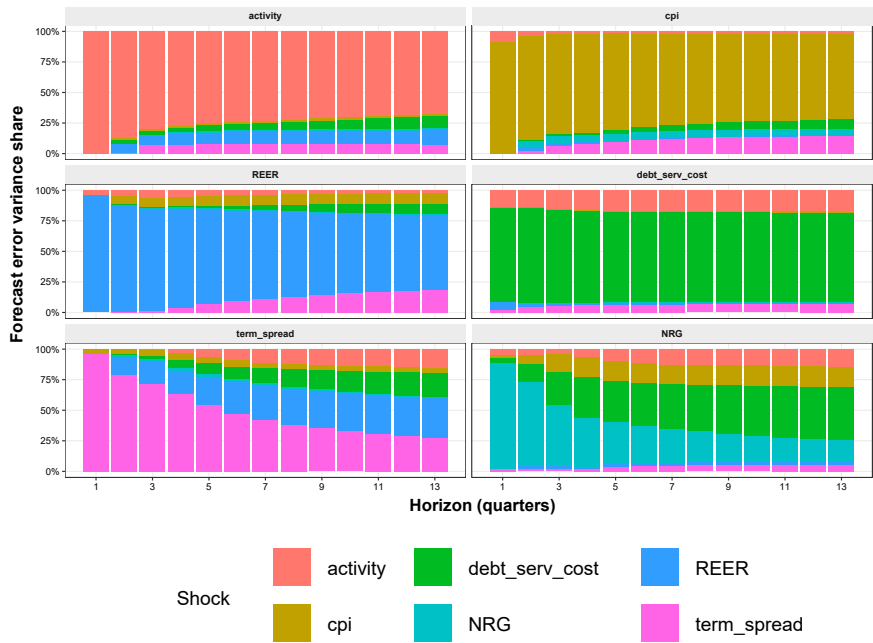


Figure C3. Baseline SVEC FEVD across all identified shocks

Source: author’s analysis.

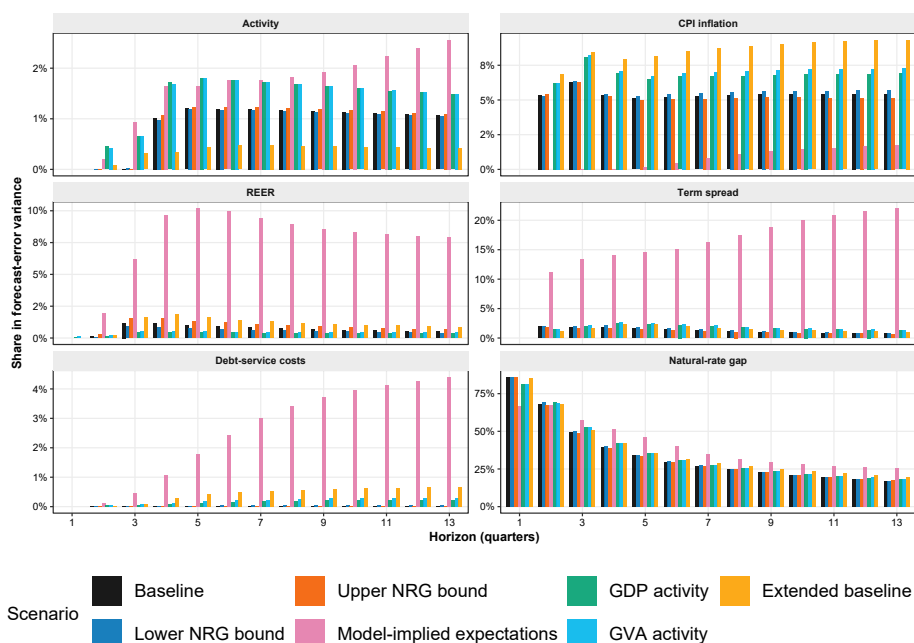


Figure C4. Robust FEVD contribution of the NRG shock across SVEC scenarios

Source: author's analysis.

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