



# DISCUSSION PAPERS

DP/125/2025

From Low to High Inflation: Regime Shifts  
and Changes in Economic Shock Transmission

by Daniel Kasabov and Militsa Slavova



BULGARIAN NATIONAL BANK

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**Abstract:** This paper investigates how economic shocks propagate under different inflation regimes in Bulgaria, focusing on the non-linear dynamics that emerge between high- and low-inflation environments. Combining theoretical insights with empirical analysis, the study explores how key features of inflation – volatility, sectoral co-movement, persistence, inflation expectations, and the wage–price nexus – vary across regimes. Using a Threshold Vector Autoregression (TVAR) model, distinct inflation regimes in Bulgaria over the past two decades are identified, and the regime-specific propagation of domestic and global shocks is assessed. The empirical analysis and model results indicate that high-inflation periods are characterized by greater volatility, stronger price synchronization across HICP groups, and a higher pass-through of cost and demand shocks to consumer prices. The magnitude of this pass-through depends on both the type and the size of the shock. Conversely, low-inflation regimes are characterized by a decoupling of wages from prices and a significantly diminished transmission of economic shocks to consumer prices, irrespective of whether such shocks stem from the domestic or external environment. These regime-specific differences in shock transmission reveal pronounced non-linearities driven by shifts in the behaviour of economic agents. This underscores the importance of macroeconomic policies that adjust to the prevailing inflationary environment, with particular emphasis on well-targeted, countercyclical fiscal measures designed to prevent the amplification of inflationary pressures, especially during periods of adverse global supply shocks.

**Keywords:** Bulgarian inflation, shock pass-through, non-linearities, inflation regimes, threshold VAR

**JEL classification:** C32, E31, E58

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# 1. Introduction

Inflation is a state-dependent phenomenon, with its characteristics evolving according to the broader macroeconomic environment. Over the past two decades, Bulgaria has experienced both low- and high-inflation periods, providing a valuable case study to explore how the macroeconomic transmission mechanism varies across these states. This paper focuses on understanding how economic shocks – both domestic and global – propagate through the economy differently depending on whether inflation is in a high- or low-inflation regime.

Traditional macroeconomic models often rely on linear approximations, assuming that the pass-through of shocks is stable across time and context. However, a growing body of literature suggests that inflation behaves in a regime-dependent manner, with firms and households responding differently to economic shocks across inflation environments (Borio et al., 2023). Such nonlinearities have important implications for interpreting inflation developments, making projections, and formulating appropriate policy responses.

In low-inflation environments, the behaviour of economic agents tends to be dominated by sector-specific and idiosyncratic factors. Firms adjust prices infrequently, workers are less focused on cost-of-living adjustments (Bobeica et al., 2019), and both consumers and producers pay less attention to inflationary signals (Sims, 2010). The “rational inattention” hypothesis posits that when inflation is subdued and stable, it becomes background noise rather than a central economic variable (Maćkowiak et al., 2023). Consequently, shocks to input costs, exchange rates, or global commodity prices often have limited and short-lived effects on headline inflation.

By contrast, high-inflation regimes are characterized by a marked shift in economic agents’ behaviour. When inflation exceeds a certain threshold, firms and workers incorporate inflation expectations more prominently in price and wage-setting decisions, making inflation more persistent and responsive to shocks. Inflation becomes salient, influencing wage negotiations, pricing strategies, and policy expectations. Firms begin to index prices more aggressively, workers demand compensatory wage increases, and the pass-through from input prices to consumer prices becomes more direct and persistent (Kumar & Wesselbaum, 2024a and 2024b). Inflation no longer reflects idiosyncratic sectoral developments but rather a broad-based inflationary pressure. The inflation process becomes more persistent and prone to feedback loops such as the classic wage-price spiral. Moreover, during high-inflation periods, expectations become more easily unanchored, amplifying the effects of shocks and undermining the stabilizing role of monetary policy (Coibion et al., 2018).

This regime-dependent behaviour is particularly relevant for Bulgaria. As a small and open economy, Bulgaria is highly sensitive to global economic developments, especially fluctuations in energy and food prices. Additionally, Bulgaria's economic structure – characterized by adverse demographic trends, a significant rise in labour market shortages since the COVID-19 pandemic, a high import content in domestically produced goods and services, and the currency board arrangement – adds further complexity to inflation dynamics. The paper empirically identifies episodes of high and low inflation using multiple quantitative approaches, including regime-switching models, and shows that the stylized facts documented in the academic literature also apply to Bulgaria.

To capture the differences in inflation behaviour across regimes, the paper explores five key features: (1) inflation volatility, (2) the degree of co-movement across HICP<sup>1</sup> groups, (3) inflation persistence, (4) the prominence of inflation expectations in firms' behaviour, and (5) the strength of the wage-price pass-through. To move beyond partial empirical-based analysis and to fully understand how shocks propagate across regimes, the study employs a Threshold Vector Autoregression (TVAR) model. This model enables the estimation of regime-specific impulse responses to various types of shocks, including global supply shocks and domestic demand shocks. Beyond shock dependence, the TVAR framework allows the pass-through (impulse response functions) to vary with the magnitude of the economic shock (size dependence).

The paper is structured as follows. Section 2 reviews existing research on inflation regimes, emphasizing the structural and behavioural mechanisms that underpin regime shifts in inflation dynamics. Section 3 presents stylized facts on inflation in Bulgaria. Drawing on descriptive evidence and existing research, it motivates the hypothesis that distinct high- and low-inflation regimes, consistent with those studied in the literature, can also be identified in Bulgaria. Section 4 presents the threshold VAR framework used to formally identify inflation regimes in Bulgaria and to examine how economic shocks propagate across them. It further details the identification scheme used to distinguish between domestic and external demand and supply shocks. Section 5 presents the empirical results derived from the TVAR estimation. Section 6 concludes.

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<sup>1</sup> HICP stands for the harmonized index of consumer prices.

## 2. Literature Review

According to the economic literature, the behaviour of economic agents – particularly in pricing and wage-setting decisions – changes markedly when the economy transitions between low- and high-inflation environments (Borio et al., 2023). This reflects the characteristics of the inflation process in both regimes. In a low-inflation environment, inflation mostly reflects sector-specific price changes that are only loosely correlated with each other. The systematic (common) component of price changes across goods and services – which corresponds to the theoretical definition of underlying inflation (Borio et al., 2021) – is limited in low-inflation environments. As a result, inflation becomes less salient in the decision-making of firms and households, leading to a muted pass-through of economic shocks to consumer prices. In addition, the notable decrease in the prominence of the common component contributes to the marked reduction in overall inflation volatility in a low-inflation environment. This decline in volatility is primarily attributed to a reduction in the co-movement (covariance) of price changes across sectors, rather than a decrease in the variability of prices within narrowly defined sectors. Equally important, wages and prices are only loosely linked, leading to subdued inflation in the labour-intensive services sector. From a theoretical perspective, this pattern of price and wage developments in a low-inflation environment can be explained by four groups of factors. First, the “rational inattention” hypothesis, introduced by Sims (2010), suggests that individuals and firms selectively process information due to cognitive and informational constraints<sup>2</sup>. In a low-inflation environment, agents are less focused on inflation signals, as the risk of significant price changes is low. This leads to delayed price adjustments and a slow response to inflationary shocks<sup>3</sup>. Second, Taylor (2000) highlights that when prices are set for multiple periods in advance due to menu costs, the lower likelihood of persistent cost- and price-related shocks in a low-inflation regime reduces firms’ incentives to adjust prices frequently. Consequently, the pass-through from costs to prices weakens, reinforcing price stickiness and muted inflationary responses. Third, the presence of downward nominal wage rigidities, as highlighted by Daly and Hobijn (2014), helps explain the flattening of the Phillips curve in low-inflation environments. The authors show that both the slope and curvature of the Phillips curve depend on the prevailing inflation rate and the extent of downward wage rigidities. At high

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<sup>2</sup> See also Maćkowiak et al. (2023). Empirical evidence about the rational inattention hypothesis can be found in Cavallo et al (2017), Coibion et al. (2020), Goodspeed (2023) and Niu and Harvey (2023).

<sup>3</sup> See also Akerlof et al. (2000), who describe price- and wage-setting behaviour at low inflation rates as “near-rational”, arguing that the costly monitoring of inflation may not be worthwhile once inflation falls below a certain threshold. In their framework, near-rationality gives rise to a long-run Phillips curve in which low, but positive, inflation is associated with higher average output levels – an effect that disappears once inflation rises sufficiently for fully rational price- and wage-setting to become optimal.

inflation, the long-run Phillips curve is nearly vertical, whereas it flattens when inflation is low. This occurs because downward rigidities prevent nominal wages from declining even when economic conditions deteriorate. Consequently, real wages remain elevated, raising labour costs for firms and potentially increasing unemployment during economic slowdowns. Fourth, the market power of firms and their ability to pass on higher costs to prices are closely linked to consumer search intensity, which itself may depend on the prevailing level of inflation. At low inflation, a large share of buyers typically observes only a single price. In this context, any economic shock significantly increases price dispersion, prompting greater consumer search activity. This heightened search reduces firms' market power, thereby limiting their ability to pass cost increases through to prices (Head et al., 2010; Bobeica et al., 2019).

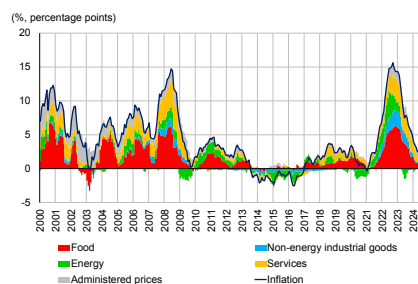
By contrast, empirical evidence indicates that once inflation surpasses a certain threshold and leaves the “rational inattention zone” a number of changes in the behaviour of agents get triggered. First and foremost, as inflation becomes a focus of attention, workers and firms will initially try to make up for the erosion of purchasing power or profit margins that they have already incurred. In addition, if inflation becomes sufficiently high and is expected to persist, they will also try to anticipate future changes in the general price level, as these will erode purchasing power and profit margins before contracts can be renegotiated. This leads to more frequent upward price adjustments, since the cost of keeping them constant increases. There is research evidence that in a high inflation environment the average length of purchase and sales contracts becomes shorter (Devereux and Yetman, 2002; Alvarez et al, 2011a, 2011b, 2019; Riggi and Tagliabracchi, 2022; Kumar and Wesselbaum, 2024a and 2024b). Moreover, in a high-inflation environment, firms may choose to preserve current profits rather than use them to absorb cost shocks, as anticipated interest rate increases induced by central banks would worsen borrowing conditions and compress future profit margins, thereby reinforcing stronger pass-through from costs to prices. In addition, heightened economic uncertainty in such environments further discourages firms from using profit margins to absorb cost increases (Bobeica et al., 2019). Under high inflation, consumers pay less attention to individual prices due to frequent price changes and heightened overall uncertainty, reducing search intensity. This allows firms to have greater market power and increase the pass-through of costs to consumer prices. As a result, inflation takes on a more prominent role in aligning the behaviour of economic agents. Therefore, during a high-inflation episode, the common component of price changes becomes more significant, firms' pricing power strengthens, wages and prices are more closely linked, and inflation becomes particularly sensitive to movements in salient prices, such as food and energy. In addition, inflation indexation clauses become more prevalent in contracts, further increasing the persistence of inflation.

### 3. Price Dynamics: Stylized Facts from Bulgaria

Data for Bulgaria suggest that inflation exhibits many of the characteristics outlined in the literature in low- and high-inflation episodes. This section presents key stylized facts for inflation developments in Bulgaria. Figure 1 depicts annual inflation in Bulgaria, with three high-inflation periods standing out: 2000–2002, 2004–2008, and 2022–2023. In all three periods, surging food and energy prices were key drivers of inflation, reflecting global commodity price developments and, in the early 2000s, the lingering effects of price liberalization and changes in administered prices. The external price pressures during the 2022–2023 episode were closely linked to pandemic-related supply chain disruptions and the economic repercussions of the war in Ukraine. Domestic factors contributing to high inflation in those periods included strong consumption growth alongside a declining unemployment rate – particularly during 2004–2008 and 2022–2023 – and positive economic expectations, especially in 2004–2008 amid high FDI inflows and EU accession. Broad-based wage increases during the periods 2004–2008 and 2022–2023, driven in part by increases in the minimum wage and public sector wages, raised labour costs and were reflected in elevated services inflation.

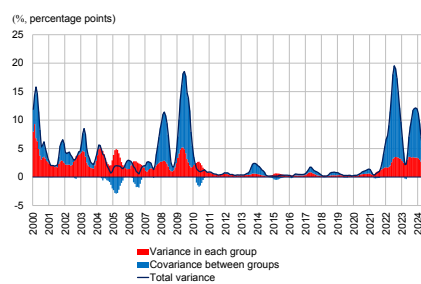
At the same time, Bulgaria experienced relatively low inflation from 2009 to 2021. Between mid-2013 and 2016, the economy entered a period of deflation, driven in part by declines in energy prices. Domestic consumption remained subdued during this time, reflecting weak credit growth, high unemployment rate, higher real interest rates and increased precautionary savings in the aftermath of the Global financial crisis. Over the full sample period, Bulgaria’s inflation process has not followed a linear path but rather a regime-dependent one, with transitions between low and high inflation closely linked to episodes of external shocks, which have been the main driver of shifts in domestic macroeconomic conditions.

Figure 1: Annual HICP inflation broken down by main groups



Source: NSI, own calculations.

Figure 2: Volatility (variance) of inflation



According to the academic literature, as inflation accelerates, sectoral price changes tend to move more closely together, indicating stronger price spillovers across sectors. These spillovers amplify inflationary pressures and increase inflation persistence, thereby reinforcing a high-inflation regime. *Figure 2* illustrates this pattern by showing the cross-group variance of annual HICP inflation, which can be decomposed into two components: the variances of individual group price changes and the covariances between them<sup>4</sup>. The results reveal that during high-inflation periods, the increase in overall variance is largely driven by higher covariances across HICP groups. In other words, inflation in such periods reflects broad-based movements in the general price level rather than isolated sectoral shocks. However, this pattern is less pronounced in the high inflation years of 2001–2004. This period exhibits distinct structural features that differentiate it from later high-inflation episodes. Specifically, it reflects a phase of economic restructuring and accelerated privatisation, characterised by high unemployment, a negative output gap and significant relative price adjustments stemming from the alignment of administratively regulated prices with production costs, as well as changes in indirect tax rates (such as the introduction of VAT on pharmaceuticals and changes in excise duties). As a result, price dynamics during 2001–2004 were driven more by one-off relative price corrections and structural adjustments than by generalised inflationary pressures. By contrast, in the post-2004 high-inflation periods, price increases were more synchronised across sectors, indicating the emergence of a genuine common inflationary component.

Another way to assess the degree of co-movement among HICP components is through principal component analysis (PCA), described in detail in Appendix 1. *Figure 3* reports the share of the variance in the monthly growth rates of 30 main HICP groups explained by the first principal component. This can be interpreted as a measure of the emergence of a “common inflation component”, indicating systemic inflationary pressures rather than isolated sector-specific shocks. The results suggest that price synchronization in Bulgaria increased markedly during 2007–2009 and again from 2021 to mid-2024, and to a lesser extent during 2002–early 2003 – periods that coincide with episodes of elevated inflation.

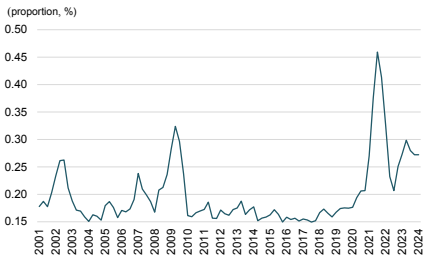
An examination of the loadings of the first principal component shows that no single HICP category dominates the component. Instead, the component reflects a broad common factor across a large number of consumption groups. The largest loadings are observed for catering services, maintenance and repair of dwellings,

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<sup>4</sup> The cross-group variance of annual HICP inflation is decomposed using the volatility of the twelve main HICP categories, as defined by Eurostat. These include processed and unprocessed food, non-energy industrial goods (durables, semi-durables and non-durables), energy components (electricity, gas, solid fuels, liquid fuels and transport fuels), and services (communication, housing, recreation, personal care, package holidays and transport). Results are robust when using a more disaggregated HICP breakdown.

personal care, clothing, and several household-related categories such as furniture, household maintenance services and household tools. At the same time, food prices also contribute relatively strongly to the first principal component. The relatively similar magnitude of the loadings across many groups suggests that the first principal component captures a common inflationary dynamic shared by a wide range of goods and services rather than price movements driven by a single dominant category.

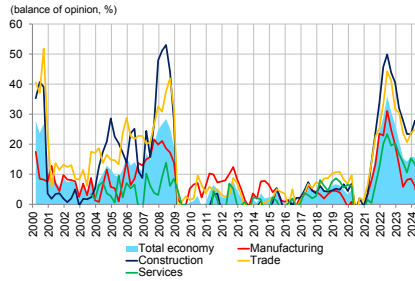
Figure 3: Degree of co-movement across HICP groups (share of variance in monthly percentage changes explained by the first principal component)



Note: The first principal component and the corresponding share of variance explained by it are estimated using an 18-month rolling window applied over the period from January 2000 to June 2024.  
Source: NSI, own calculations.

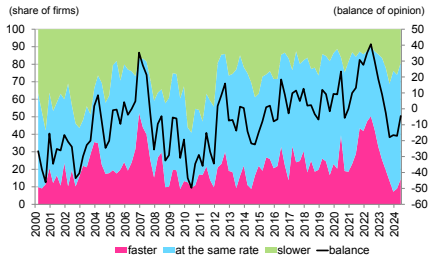
A complementary indicator of regime shifts is the behaviour of inflation expectations, which reflect changes in inflation psychology. *Figure 4* shows that during the identified high-inflation episodes, there was a broad-based rise in three-month selling price expectations among firms across the four main economic sectors – manufacturing, construction, trade, and services – based on regular business surveys conducted by the NSI. In the trade sector, the surveys also include a question about the magnitude of expected price increases. *Figure 5* reveals that during the 2004–2008 and 2022–2023 high-inflation episodes, not only did the share of retail trade firms anticipating price increases rise sharply, but so did the share of firms reporting expectations of faster price increases.

Figure 4: Inflation expectations of firms



Note: Data is seasonally adjusted by the author.  
Source: NSI, own calculations.

Figure 5: Magnitude of price increases in retail trade



Kasabov (2024) highlights a non-linear cost pass-through to consumer prices during high-inflation periods combined with strong domestic demand, illustrating a steepening of the Phillips curve in Bulgaria. The author estimates the following reduced-form Phillips curve model:

$$\pi_t = \mu_t + \rho_t \pi_{t-1} + \theta_t \pi_t^e + \beta_{1t} \Delta ulc_t + \beta_{2t} \pi_t^{imp} + \beta_{3t} x_t + e^{\frac{h_t}{2}} \varepsilon_t$$

where  $\pi_t$  represents inflation (quarterly percent change in the headline consumer price index) at time  $t$ ,  $\mu_t$  is a time-varying intercept parameter,  $\pi_t^e$  denotes expected inflation of firms at time  $t$ ,  $\Delta ulc_t$  is the quarterly change in nominal unit labor costs at time  $t$ ,  $\pi_t^{imp}$  is import price inflation,  $x_t$  is a measure of slack (the unemployment rate) and  $\varepsilon_t$  is a shock term with a time-varying volatility function  $e^{\frac{h_t}{2}}$ .

Figure 6 and Figure 7 present results from the estimated Phillips curve. Figure 6 shows a gradual decline in the unit labour cost (ULC) coefficient after 2008, indicating a flattening of the inflation–ULC relationship. Between 2002 and 2008, a 1% increase in ULC raised consumer prices by about 0.11% in the same quarter (0.12% by late 2007). Following the Global Financial Crisis (2009–mid-2013), the short-term impact dropped to around 0.09%–0.10%. From 2022 onward, a 1% ULC increase corresponded to a 0.16% rise in consumer prices within the same quarter. The results from the model imply that amid significant supply shocks, uncertainty, and high inflation expectations as well as in the context of strong

<sup>5</sup> The model is estimated using seasonally adjusted series, with the unemployment rate expressed in levels and inflation expectations in first differences. The quarterly growth rates of the remaining series are calculated using log differences. The specification that is employed stands out with respect to standard single-equation Phillips curve models in two directions. First, it allows the coefficients on explanatory variables to be time-dependent, which implies that the structural relationships between inflation and its drivers might change over time. Second, the variance of the unobserved shocks is also allowed to change over time. Bayesian methods are employed for estimating the model parameters.

private consumption, price adjustments over the 2022–2023 period became more frequent and immediate as the cost of maintaining stable prices rose.

Figure 6: Estimates of the impact multiplier (contemporaneous effect) of ULC,  $\beta_{1t}$

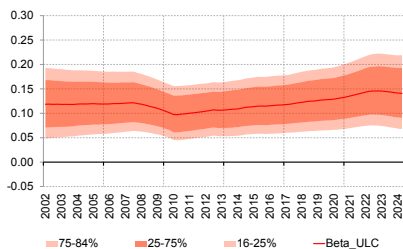
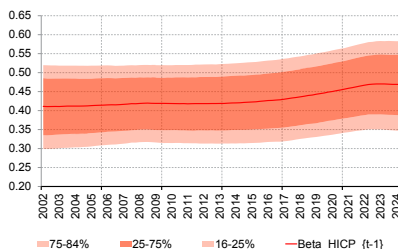


Figure 7: Estimates of inflation persistence (coefficient on lagged inflation),  $\rho_t$

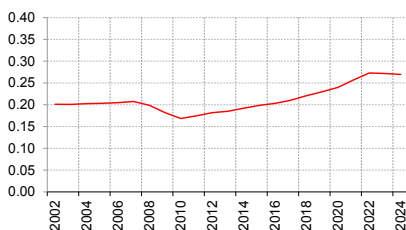


Note: Data is seasonally adjusted by the author.

Source: NSI, own calculations.

The rise in the slope of the Phillips curve after the COVID-19 pandemic was also accompanied by higher inflation persistence, as indicated by the coefficient on lagged inflation ( $\rho_t$ ) in the Phillips curve model (see *Figure 7*). This, in turn, implied a larger long-run pass-through of labour costs to consumer prices<sup>6</sup>, as shown in *Figure 8*.

Figure 8: Labour cost pass-through in the long-run



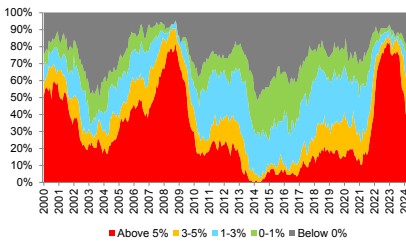
Source: NSI, own calculations.

<sup>6</sup> Regarding the estimated Phillips curve model, the long-run (cumulative) impact of labour costs on prices depends on both the slope of the Phillips curve ( $\beta_{1t}$ ) and inflation persistence ( $\rho_t$ ). Mathematically, the cumulative increase in the consumer price index after  $n$  periods following a 1% increase in ULC can be estimated as  $\beta_{1t} * (1 + \rho_t + \rho_t^2 + \dots + \rho_t^{n-1})$ .

In the absence of individual firm-level price data, the frequency and magnitude of price adjustments can be assessed indirectly using a diffusion index that captures the share of HICP subcomponents experiencing price changes within pre-defined intervals. *Figure 9* illustrates that during the high-inflation episodes of 2000–2001, 2007–mid-2009, and 2022–early 2023, more than half of HICP groups registered annual price increases of over 5%. These episodes stand in sharp contrast to the low-inflation environment observed between mid-2013 and 2016, when the majority of price changes were confined to the 0–1% and 1–3% ranges. During this period, the share of HICP components exhibiting outright deflation fluctuated between 30% and 50%, indicating that firms were likely facing a pronounced negative demand shock.

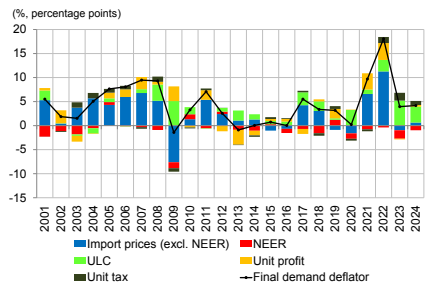
Evidence from firm-level survey data supports this interpretation. Kasabov and Paskaleva (2018) report that between 2009 and 2013, firms’ pricing and cost-setting decisions were primarily driven by changes in macroeconomic conditions and the intensity of negative shocks – specifically, shocks related to external financing, access to usual suppliers, and clients’ ability to pay. Firms simultaneously hit by these shocks were significantly more likely to reduce prices and less likely to increase them compared to firms not affected. Similarly, cost adjustments were more substantial among adversely affected firms: those experiencing negative shocks in 2009 were less inclined to raise costs, indicating efforts to optimize production structures following the onset of the Global financial crisis.

Figure 9: Diffusion index



Source: NSI, own calculations.

Figure 10: Breakdown of the final demand deflator



Using national accounts data, firms’ mark-up behaviour can be inferred from the dynamics of the final demand deflator. *Figure 10* decomposes the growth rate of the final demand deflator into the contributions of import prices, the nominal effective exchange rate (NEER), unit labour costs (ULC), profit per unit of output, and taxes (net of subsidies) per unit of output. The results indicate

that during high-inflation episodes, such as 2004–2008 and 2022–2023, firms in Bulgaria were able to sustain higher unit profits, suggesting an increase in mark-ups and a stronger pass-through of production cost pressures to consumer prices.

In summary, inflation dynamics in Bulgaria display clear state dependence, with high- and low-inflation periods exhibiting distinct patterns influenced by both external and domestic factors. High-inflation episodes are characterised by broad-based price increases, stronger sectoral spillovers, heightened inflation persistence, and an amplified pass-through of labour costs to consumer prices. Since the emergence and persistence of high-inflation regimes depend on both sector-specific shocks and their propagation across the economy, the next section employs a TVAR framework to examine how the transmission of macroeconomic shocks differs across inflation regimes.

#### 4. Methodology

To examine the state-dependent transmission of shocks across different inflation regimes, this study employs the TVAR model introduced by Alessandri and Mumtaz (2017). The general TVAR model is defined as follows:

$$Y_t = (c_1 + \sum_{j=1}^p B_{1,j} Y_{j,t-j} + \Omega_1^{1/2} v_t) S_t + (c_2 + \sum_{j=1}^p B_{2,j} Y_{j,t-j} + \Omega_2^{1/2} v_t) (1 - S_t) \quad (1)$$

where  $Y_t$  is the matrix of endogenous variables and  $v_t \sim N(0,1)$ . The state function  $S_t$  takes the value of 1 if  $Z_{t-d} \leq Z^*$  and 0 otherwise, defining two distinct regimes based on the level of the threshold variable  $Z_{t-d}$  relative to an unobserved threshold  $Z^*$ . In this model, the threshold variable is assumed to be the  $d^{\text{th}}$  lag of annual HICP inflation, with  $d$  treated as an unknown parameter. Consequently, both the delay  $d$  and the threshold  $Z^*$  are estimated. The model incorporates two sets of parameters, one for each regime, defined by  $\{c_s, B_{s,j}, \Omega_s\}$  for  $s = \{1,2\}$ . Here  $c_1$  and  $c_2$  are vectors of constants,  $B_1$  and  $B_2$  denote the coefficient matrices of the TVAR model,  $\Omega_1$  and  $\Omega_2$  represent the covariance matrices of the residuals.

This specification of equation (1) allows both the transmission mechanism and the magnitude of economic shocks to vary across high- and low-inflation regimes. The lag order of the model,  $P$ , is set to 2 based on standard lag-length selection criteria.

In this study, the TVAR model is applied to examine the impact of domestic and global shocks on HICP inflation in Bulgaria. The baseline specification of the TVAR model includes four endogenous variables: real GDP, consumer prices (HICP), import prices, and real exports. Therefore, the vector of endogenous variables is defined as  $Y_t = \{gdp_t, hicp_t, import\ prices_t, exports_t\}$ . These variables enable an analysis of how demand and supply shocks, originating

from both global and domestic sources, affect HICP inflation. The inclusion of import and consumer price variables in the model is intended to capture the propagation of price pressures through the supply chain and to highlight potential differences between inflation regimes<sup>7</sup>.

The data are quarterly, spanning from the first quarter of 2000 to the first quarter of 2024. All series in equation (1) are seasonally adjusted and included in the model as log differences. As mentioned earlier, the threshold variable  $Z_{t-d}$  is defined by the year-on-year growth rate of HICP. This transformation is selected due to the lower volatility of annual inflation rates compared to annualized quarterly growth rates, as well as the common practice of reporting inflation on an annual basis. High quarterly inflation in individual quarters may reflect data volatility rather than genuine inflationary trends, whereas high year-on-year inflation rates are more likely to indicate a true strengthening of price pressures in the economy. All data are sourced from the National Statistical Institute, with the seasonal adjustment of the HICP series performed by the authors. The data for real GDP, real exports, and import prices are obtained from national accounts statistics.

The relatively short sample period and the complexity of the model justify the parsimonious structure of the TVAR model, which includes only four endogenous variables. The estimation of the TVAR model closely follows the approach outlined by Alessandri and Mumtaz (2017). Model parameters are estimated using Bayesian methods, with a natural conjugate prior imposed on the TVAR parameters in both regimes. The Gibbs sampler, as introduced by Chen and Lee (1995), is employed to simulate the posterior distribution of the unknown parameters. Detailed information on the tightness of the priors and the starting parameter values required for the initialization of the Gibbs sampling algorithm is provided in Appendix 2.

The analysis of economic shock pass-through in low and high-inflation regimes is conducted using structural impulse response analysis. Following Alessandri and Mumtaz (2017), generalized impulse responses, as proposed by Koop et al. (1996), are employed. Unlike traditional impulse responses, generalized impulse responses are more appropriate for non-linear models as they account for both history-dependence and shock-dependence, considering the size and sign of

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<sup>7</sup> One variable that could potentially be included in the model is wage growth, given its relevance for domestic cost pressures and inflation dynamics. However, introducing wages as an additional endogenous variable would significantly increase the dimensionality of the threshold VAR system and the number of parameters to be estimated, which may reduce the precision of the estimates given the available sample size. Moreover, wage dynamics partly reflect the same domestic demand conditions captured by the demand variable already included in the model, which may create additional identification challenges when distinguishing between demand- and cost-driven shocks. For these reasons, the baseline specification focuses on a more parsimonious set of variables while allowing the effects of wage developments to be reflected indirectly through domestic demand conditions.

the shock. These generalized impulse responses are computed using Monte Carlo integration. Ultimately, the impulse response functions (IRFs) for the two regimes  $S = \{0,1\}$  are derived as the differences between two conditional expectations: one based on simulating the model with a structural shock  $\mu$  and the other where all shocks are set to zero. Specifically:

$$IRF_t^S = E(Y_{t+k} | \psi_t, Y_{t-1}^S, \mu) - E(Y_{t+k} | \psi_t, Y_{t-1}^S) \quad (2)$$

where  $\psi_t$  includes the parameters and hyperparameters of the TVAR model,  $Y_{t-1}^S$  denotes the regime-specific history, and  $k$  refers to the horizon considered.

The impulse responses are conditioned on observations within each regime. Specifically, the impulse response for each regime is computed as the average of impulse responses obtained by simulating the model from all starting values within that regime. As mentioned, generalized impulse responses are history-dependent and influenced by the starting values within each regime. For example, in the high-inflation regime, the impact of a shock may vary depending on whether it occurs when inflation is significantly above or just slightly above the threshold. While impulse responses may differ within each regime, the goal of calculating them conditional on the observations of the respective inflation regimes is to capture the average or most typical behaviour of the economy in each regime (see also Alessandri and Mumtaz (2019)).

Additionally, generalized impulse responses account for the possibility of endogenous regime switches. This implies that the economy can transition between regimes over the simulation horizon, as both the evolution of endogenous variables and the parameters of the TVAR model are considered. In this paper, for instance, inflation is both an endogenous variable and the threshold variable of the TVAR model. Consequently, a shock affecting future price developments can lead to a change in the regime over the simulation horizon.

As previously mentioned, the baseline VAR includes two lags of the following four variables: Bulgarian real GDP, HICP, import deflator, and real exports. The study identifies four types of structural economic shocks: a global supply (commodity price) shock, a global demand shock, a domestic aggregate demand shock, and a domestic aggregate supply shock. In the empirical analysis, the focus is on the global supply and domestic demand shocks, as these two types of shocks drive the two main types of inflation: demand-pull and cost-push inflation.

The restrictions imposed to identify the four shocks are outlined in *Table 1* and are based on several widely accepted principles. Demand and supply shocks are linked to positive and negative correlations between real activity and prices. In the context of a small and open economy, domestic variables are assumed not to influence global developments, which helps to distinguish local economic shocks

from those generated abroad. As shown in studies by Comunale and Kunovac (2017), Forbes et al. (2018), and Kronborg (2021), a combination of zero short-run and sign restrictions is required to accommodate the small open economy assumption. The orthogonalization process, which involves applying zero and sign restrictions, as described by Arias et al. (2018), is implemented in this analysis.

Table 1: Identification pattern based on zero and sign restrictions

Shock Variable	Domestic demand	Domestic supply	Global demand	Global supply (foreign prices)
Real output	+	-	+	-
Inflation	+	+	+	+
Import prices	•	•	•	+
Exports	0	0	+	•

Notes: "\*" = unconstrained, "+" = positive sign, "-" = negative sign, "0" = zero restriction. All restrictions are imposed on impact.

Given that Bulgaria is characterized as a small open economy, the baseline specification assumes that domestic aggregate demand and supply shocks cannot influence foreign demand in the short run. Therefore, zero restrictions are imposed on exports, preventing domestic demand and domestic supply shocks from affecting the exports of goods and services. This restriction is applied only on impact, as the propagation of domestic shocks throughout the economy and their effects on prices may eventually alter the economy's competitive position, thus influencing export activity in the medium run.

In summary, for a domestic demand shock, the model imposes standard restrictions, requiring that both real output (GDP) and HICP increase in response to a positive shock. For an inflationary domestic supply shock, the model requires that HICP increase while output declines. In the case of a positive global demand shock, exports, domestic GDP, and consumer prices all increase. Finally, for an inflationary global supply shock, it is assumed that import prices rise, domestic HICP increases, and output decreases.

## 5. Empirical Results

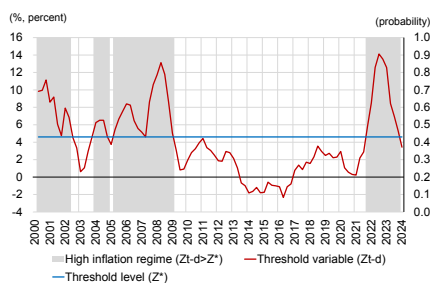
This section presents the empirical results on the pass-through of economic shocks to consumer prices in Bulgaria under high- and low- inflation regimes. The analysis focuses on global supply and domestic demand shocks, as these are the primary drivers of cost-push and demand-pull inflation, respectively.

Figure 11 presents annual HICP inflation together with the two inflation regimes identified by the threshold VAR model estimated over the period from the first

quarter of 2000 to the first quarter of 2024. The shaded areas represent periods where annual inflation exceeds the estimated threshold of 4.6 percent, which is obtained endogenously from the Bayesian estimation of the threshold VAR model described in Section 4. Inflation below this threshold is classified as the “low inflation” regime, while periods with year-on-year HICP growth exceeding this threshold are labelled as “high inflation.”<sup>8</sup> According to the estimates, the Bulgarian economy was in a “high inflation” regime from the first quarter of 2000 to mid-2002, from the first quarter of 2004 to the first quarter of 2009 (excluding Q1 2005), and from the end of 2021 to the fourth quarter of 2023. These periods of high inflation closely align with the narrative and descriptive evidence presented in Section 3.

From an economic perspective, the estimated threshold appears plausible. An inflation rate of around 4–5 percent is sufficiently high to make inflation more salient in the decision-making of firms and households, thereby triggering behavioural adjustments in price- and wage-setting. This interpretation is consistent with the literature on inflation regimes (e.g. Borio et al., 2023), which suggests that once inflation rises beyond a moderate range, economic agents begin to incorporate inflation expectations more actively into their decisions. In the Bulgarian case, the threshold also aligns well with periods of stronger inflation synchronisation across HICP components and increased pass-through of demand and cost shocks to consumer prices identified in the empirical analysis.

Figure 11: HICP year-on-year percentage change and estimated inflation regimes



Note: The shaded areas represent the periods where annual HICP inflation is above the threshold.

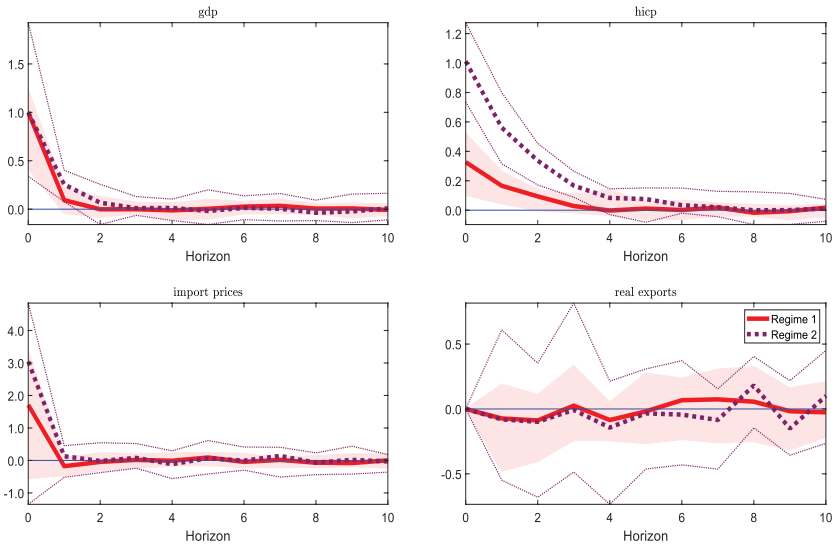
Figure 12 illustrates the estimated impact of a positive domestic demand shock on the four variables included in the baseline model across the two inflation regimes.

<sup>8</sup> For clarity, in the figures that follow, the regime in which year-on-year HICP inflation is below the estimated threshold of 4.6 percent is referred to as Regime 1, while the regime in which inflation exceeds this threshold is referred to as Regime 2. For ease of exposition, these regimes are also described as low- and high-inflation regimes, respectively. The terminology is used purely as a convenient shorthand for the two states identified by the threshold VAR model and does not correspond to broader classifications of inflation levels in the macroeconomic literature.

The size of the shock is one standard deviation under each regime. To ensure comparability of the results, the responses of all variables to the one standard deviation demand shocks are normalized so that real output increases by one percent in the first quarter of the shock. The non-normalised impulse responses are reported in Appendix 3.<sup>9</sup> The figure presents the median impulse responses in log differences, along with the corresponding 68% credibility intervals for both regimes. The dashed lines represent responses in the high inflation regime, while the red lines and shading depict those in the low inflation regime.

Figure 12: Impulse responses to a demand shock of **one standard deviation** in each regime (normalised to a one percent increase in real GDP on impact in each regime)

(log differences)



Note: The x-axis refers to the horizon of the impulse responses in quarters.

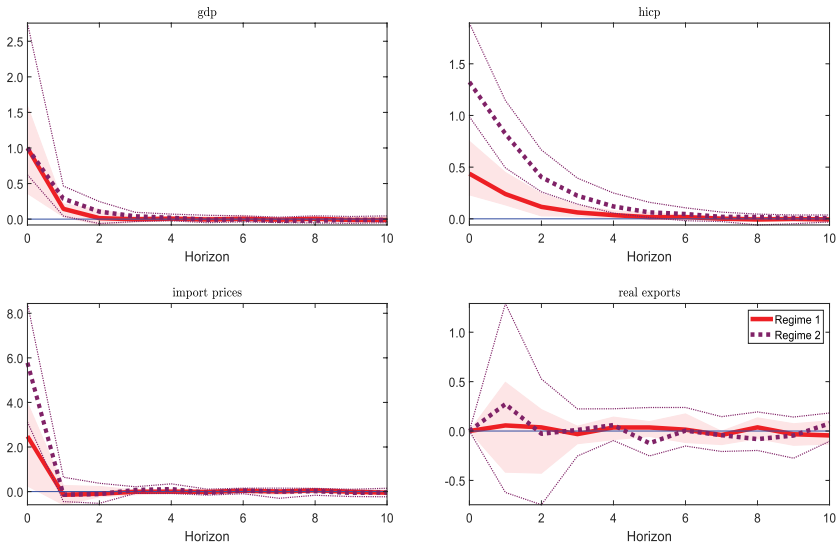
According to the model, a positive domestic demand shock leads to an increase in GDP and HICP in both regimes, consistent with the imposed first-quarter sign restrictions, while exports remain unaffected. In the high-inflation regime, median simulation estimates suggest that a 1% rise in GDP results in an HICP

<sup>9</sup> The non-normalised impulse response results indicate that the four macroeconomic variables included in the TVAR model exhibit markedly different standard deviations across regimes. In particular, GDP and import prices display significantly higher volatility under the low-inflation regime. This can be attributed to the fact that this regime encompasses the post-Global Financial Crisis and COVID-19 pandemic periods, during which sharp declines in global commodity prices and GDP were observed, contributing to the higher standard deviation of both variables.

increase of about 1%, compared with just under 0.4% in the low-inflation environment. Although no restrictions are imposed on import prices, they tend to increase modestly in the short term in both regimes, though this response is not statistically significant. We can therefore conclude that a typical domestic demand shock has an inflationary impact roughly two and a half times larger in the high-inflation regime than in the low-inflation regime.

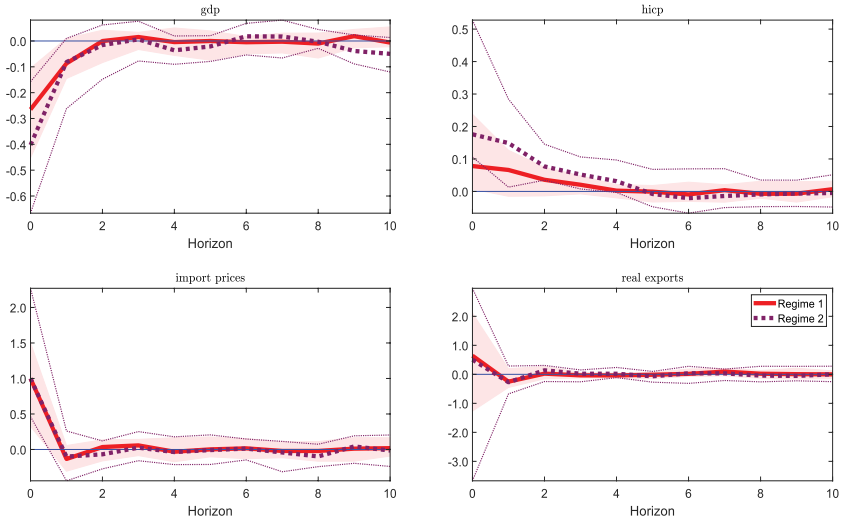
In addition to demonstrating state dependence, the baseline TVAR model suggests that the impact of a demand shock varies with its size, but only for shocks of at least three standard deviations (*Figure 13*). Following such a large shock, a 1% increase in real output leads to an average rise in HICP of around 1.3% on impact, compared to around 1% for smaller shocks. Large shocks also produce a statistically significant positive response in import prices, suggesting firms may seek to raise mark-ups. However, there is no evidence from the TVAR model that stronger domestic demand shocks have more persistent effects than smaller ones.

Figure 13: *Impulse responses to a demand shock of three standard deviations in each regime (normalised to a one percent increase in real GDP on impact in each regime)*  
(log differences)



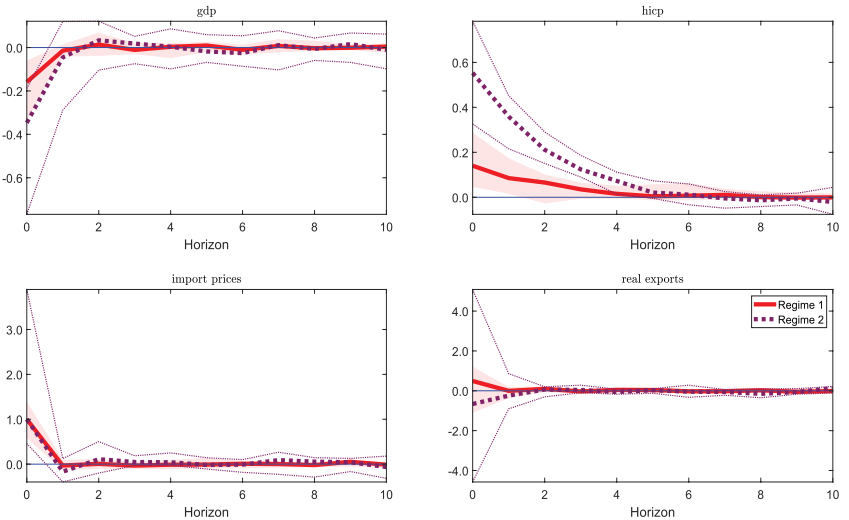
Note: The x-axis refers to the horizon of the impulse responses in quarters.

Figure 14: *Impulse responses to a global supply (foreign price) shock of one standard deviation in each regime (normalised to a one percent increase in import prices on impact in each regime)* (log differences)



Note: The x-axis refers to the horizon of the impulse responses in quarters.

Figure 15: *Impulse responses to a global supply (foreign price) shock of two standard deviations in each regime (normalised to a one percent increase in import prices on impact in each regime)* (log differences)



Note: The x-axis refers to the horizon of the impulse responses in quarters.

*Figure 14* and *Figure 15* illustrate the estimated effects of an inflationary global supply shock of one and two standard deviations on the four variables of the baseline model across the two inflation regimes. The generalized impulse responses do not point to strong disproportional effects for small global supply (foreign price) shocks in either regime. This implies that when shocks are small firms may opt to absorb them, at least partially, through profits margins. However, as the shock size increases to two or more standard deviations, there is evidence of a stronger and statistically significant pass-through from foreign to consumer prices during high inflation periods. In other words, the impact of foreign price shocks becomes increasingly nonlinear as their magnitude grows.

Overall, the empirical results confirm that inflation dynamics in Bulgaria exhibit clear state dependence, with the pass-through of domestic demand and global supply shocks to consumer prices differing significantly between high- and low-inflation regimes. Domestic demand shocks generate substantially larger inflationary effects during high inflation periods, and this impact intensifies further for shocks exceeding three standard deviations. Similarly, while small global supply shocks have limited and roughly proportional effects, large foreign price shocks lead to markedly stronger and nonlinear pass-through under high inflation. These findings underscore the importance of accounting for the prevailing regime, as well as the type and magnitude of shocks, when assessing inflationary pressures.

## 6. Conclusion

The empirical findings of this study underscore the necessity for a state-contingent understanding of inflation dynamics that explicitly accounts for regime dependence and shock heterogeneity. The results point to statistically significant differences in inflation behaviour across regimes. High-inflation phases in Bulgaria are characterised by elevated volatility, stronger co-movement of sectoral price changes, increased persistence, greater salience in price- and wage-setting decisions, and a tighter wage–price linkage. These features make inflation more difficult to stabilise and more sensitive to shocks – particularly those stemming from global supply disturbances or strong domestic demand pressures.

The distinct regime-specific patterns imply that the effectiveness of policy responses depends critically on correctly identifying the prevailing inflation environment and the magnitude of shocks hitting the economy. The evidence of nonlinear pass-through of both domestic demand and global supply shocks during high-inflation regimes suggests that conventional policy frameworks – typically built on assumptions of linearity and stability – may be insufficient. Policymakers should therefore employ flexible, state-dependent strategies calibrated not only to the level of inflation but also to the intensity and nature of economic disturbances. Such

an approach can strengthen policy credibility, anchor inflation expectations more effectively, and improve macroeconomic outcomes in Bulgaria.

Building on these insights, the study has three major implications for policymakers. First, given the constraints of the currency board arrangement and the limited monetary policy toolkit, policymakers can adopt a more flexible approach during low-inflation periods, tolerating moderate fluctuations in inflation without the need to recalibrate fiscal policy or adjust monetary instruments such as the minimum required reserve rate. Considering the small size and openness of the Bulgarian economy, fiscal authorities may implement policies with limited regard to price developments when inflation is low, as price and cost shocks tend to have short-lived effects.

Second, in a high-inflation context, characterized by large demand or cost-push shocks, it is essential to act pre-emptively when there is a heightened risk of transitioning into a high-inflation regime. Policymakers should adopt countercyclical fiscal and monetary measures that avoid fuelling domestic demand or raising production costs, given the highly nonlinear behaviour of inflation in such environments. When discretionary fiscal interventions are necessary to address the social consequences of shocks and cost-of-living crises, these measures should be carefully targeted and designed to remain neutral with respect to aggregate demand.

Third, the observed return of macroeconomic volatility following the COVID-19 pandemic underscores the need for central banks – both in Bulgaria and internationally – to adopt a preemptive and cautious stance, given the inherent delays in monetary policy transmission. The results for Bulgaria align with evidence from other economies showing that larger economic shocks tend to have disproportionately greater consequences during high-inflation regimes, irrespective of shock origin. This insight will be particularly relevant for Bulgaria as it prepares to join the euro area in 2026, where monetary policy will be conducted at the eurozone level and timely, forward-looking policy responses will remain essential for maintaining price stability.

The findings of the paper, including the estimated value of the threshold parameter and the magnitude of the impulse responses, should not be interpreted as definitive. The model results are conditional on the specific model specification, the estimation sample, the selected set of variables, and the identification strategy used to distinguish between supply and demand shocks. While the threshold VAR framework provides a useful tool for analysing regime-dependent dynamics, alternative model specifications or different identification assumptions could potentially lead to somewhat different quantitative estimates. Nevertheless, the results appear robust to reasonable variations in model settings and should therefore be interpreted as indicative evidence regarding the changing transmission of economic shocks to consumer prices in Bulgaria rather than as definitive estimates.

## Appendix 1: Principal Component Analysis for the co-movement of Inflation Subcomponents Across Time

To assess price co-movement in Bulgaria, Principal Component Analysis (PCA) is applied. PCA is a dimensionality reduction technique that condenses information from a large set of variables into a smaller number of unobserved components that capture the dominant patterns of variation. The procedure is based on the variance–covariance matrix of the data. The variance of each variable reflects its individual informational contribution, while the covariance terms capture the extent to which fluctuations across variables move together. When the covariance matrix is diagonal, each variable contributes distinct information; in contrast, sizeable off-diagonal elements indicate common underlying movements. PCA transforms the original variables into a new set of orthogonal components by identifying eigenvectors that diagonalize the covariance matrix, thereby isolating the main common source(s) of variation.

The analysis uses monthly data for 30 major HICP subcomponents spanning the period January 2000 to June 2024. A larger initial set of HICP categories was considered, but several were excluded due to shorter data availability, which would have constrained the PCA sample window. The 30 retained subcomponents are:

- 1.1 Food,
- 1.2 Non-alcoholic beverages,
- 2.1 Alcoholic beverages,
- 3.1 Clothing,
- 3.2 Footwear,
- 4.1 Actual rentals for housing,
- 4.3 Maintenance and repair of the dwelling,
- 4.4 Water supply and miscellaneous services relating to the dwelling,
- 4.5 Electricity, gas and other fuels,
- 5.1 Furniture and furnishings, carpets and other floor coverings,
- 5.2 Household textiles,
- 5.3 Household appliances,
- 5.4 Glassware, tableware and household utensils,
- 5.5 Tools and equipment for house and garden,
- 5.6 Goods and services for routine household maintenance,
- 6.1 Medical products, appliances and equipment,
- 6.2 Out-patient services,
- 7.1 Purchase of vehicles,

- 7.2 Operation of personal transport equipment,
- 7.3 Transport services,
- 8.2 Telephone and telefax equipment,
- 9.1 Audio-visual, photographic and information processing equipment,
- 9.3 Other recreational items and equipment, gardens and pets,
- 9.4 Recreational and cultural services,
- 9.5 Newspapers, books and stationery,
- 11.1 Catering services,
- 11.2 Accommodation services,
- 12.1 Personal care,
- 12.3 Personal effects n.e.c.,
- 12.7 Other services n.e.c.

A rolling-window PCA is implemented to trace the evolution of price co-movement over time and across inflation regimes, using an 18-month window. To meet PCA requirements, all series are first seasonally adjusted, then transformed into stationary dlog changes and standardized. Inspection of the eigenvalue structure indicates that retaining only the first principal component is appropriate. In this setting, the first principal component represents the dominant common factor driving price changes across HICP categories and serves as the primary indicator of co-movement.

## Appendix 2: Set-up of the Threshold VAR

The estimation of the TVAR model follows the approach applied by Alessandri and Mumtaz (2017). As described in Alessandri and Mumtaz (2017), a natural conjugate prior is imposed on the TVAR parameters of equation (1) via dummy observations, as proposed by Sims and Zha (1998) and Banbura et al. (2010):

$$Y_{D,1} = \begin{pmatrix} \frac{\text{diag}(y_1 \sigma_1 \dots y_N \sigma_N)}{\tau} \\ 0_{N(P-1) \times N} \\ \dots \\ \text{diag}(\sigma_1 \dots \sigma_N) \\ 0_{1 \times N} \end{pmatrix}, \text{ and}$$

$$X_{D,1} = \begin{pmatrix} \frac{I_p \otimes \text{diag}(\sigma_1 \dots \sigma_N)}{\tau} & 0_{NP \times 1} \\ 0_{N \times NP} & 0_{N \times 1} \\ \dots & \dots \\ 0_{1 \times NP} & c \end{pmatrix}$$

where  $\gamma_1 \dots \gamma_N$  represent the prior means for the coefficients on the first lag,  $\sigma_1 \dots \sigma_N$  denote the scaling factors,  $\tau$  is the tightness of the prior on the TVAR coefficients,  $c$  is the tightness of the prior on the constant term, and  $J_p = \text{diag}(1, 2, \dots, P)$ . The prior means are the OLS estimates of the coefficients from an AR(1) regression estimated for each endogenous variable using a training sample<sup>10</sup>, while the scaling factors correspond to the standard deviation of the error terms from these AR(1) regressions. The tightness of the prior on the TVAR coefficients is set to  $\tau = 0.1$ , and a flat prior is chosen for the constant term, setting  $c = 1/10000$ .

The choice of a natural conjugate prior follows the approach commonly used in Bayesian VAR models with relatively short samples, as it allows for efficient posterior simulation and facilitates the implementation of the Gibbs sampler. In addition, this prior structure introduces shrinkage toward a simple autoregressive process, which helps stabilize estimation in systems with several endogenous variables and limited observations. This specification follows the empirical strategy adopted by Alessandri and Mumtaz (2017) and related applications in the literature. Given the relatively small sample size and the focus on regime-dependent dynamics, the use of this standard prior structure ensures comparability with existing studies while limiting the risk of over-parameterization.

A prior is also imposed on the sum of the lagged dependent variables by adding the following dummy observations:

$$Y_{D,2} = \left( \frac{\text{diag}(\gamma_1 \mu_1 \dots \gamma_N \mu_N)}{\lambda} \right) \text{ and} \\ X_{D,2} = \left( \frac{(\mathbf{1}_{1 \times P}) \otimes \text{diag}(\gamma_1 \mu_1 \dots \gamma_N \mu_N)}{\lambda} \quad \mathbf{0}_{N \times 1} \right)$$

where  $\mu_i$  is the sample mean of the endogenous variables calculated from the training sample, and the tightness of the sum of coefficients prior is chosen as  $\lambda = 10\tau$ , following Alessandri and Mumtaz (2017) and Banbura et al (2010). The priors on the TVAR parameters are set to be identical for the two regimes, ensuring that any differences between regimes are data-driven rather than prior-induced.

<sup>10</sup> The training sample consists of the initial observations of the dataset and is used solely to calibrate the prior distribution. The same data are subsequently included in the estimation of the TVAR model. The estimated AR(1) coefficients provide a simple characterization of the persistence of each variable and serve as prior means for the VAR coefficients, while the standard deviations of the residuals from these regressions are used as scaling factors that determine the prior variance of the coefficients in each equation. We experimented with alternative values of the tightness parameters, but the results proved robust and the baseline specification follows the parameterization used in Alessandri and Mumtaz (2017).

A flat prior is chosen for the delay parameter  $d$ , with a maximum delay set to 2. Finally, a normal prior is applied to the threshold:

$$z^* \sim N(\bar{z}, \bar{v}),$$

where  $\bar{z} = \frac{1}{T} \sum_{t=1}^T z_t$  and  $\bar{v} = 10$ , which is relatively loose considering the scale of HICP inflation.

The Gibbs sampling algorithm of Chen and Lee (1995) is used to estimate the TVAR model. To that aim, the data are split into regime-specific observations. Given an initial value for  $z^*$  and  $d$ , the conditional posterior distribution for the VAR parameters  $B = \text{vec}([c, B_1; B_2 \dots; B_p])$  and  $\Omega$  in the two regimes follows a standard form given the applied natural conjugate prior:

$$G(B|\Omega) \sim N(B^*, \Omega \otimes (X^{*'} X^*)^{-1}),$$

$$G(\Omega|B) \sim IW(S^*, T^*),$$

where the posterior means are

$$B^* = (X^{*'} X^*)^{-1} (X^{*'} Y^*) \quad \text{and} \quad S^* = (Y^* - X^* \tilde{B})' (Y^* - X^* \tilde{B}).$$

Here,  $Y^* = [Y; Y_{D,1}; Y_{D,2}]$  and  $X^* = [X; X_{D,1}; X_{D,2}]$  and  $\tilde{B}$  is the draw of the VAR coefficient  $B$  reshaped to conform with  $X^*$ . The parameter  $T^*$  represents the number of rows of  $Y^*$  (see Alessandri and Mumtaz (2017)).

Given a draw for the VAR parameters and a value for  $d$ , the threshold value  $z^*$  is drawn from its non-standard posterior in a random walk Metropolis Hastings step. A candidate value of  $z_{new}^*$  is drawn as:

$$z_{new}^* = z_{old}^* + \psi^{1/2} \epsilon, \quad \epsilon \sim N(0, 1).$$

The acceptance probability is given by:

$$\frac{f(Y_t | z_{new}^*, \Xi)}{f(Y_t | z_{old}^*, \Xi)},$$

where  $f(\cdot)$  represents the posterior density and  $\Xi$  includes all other parameters in the model. The scaling factor  $\psi$  is chosen such that the acceptance rate lies between 20% and 40%.

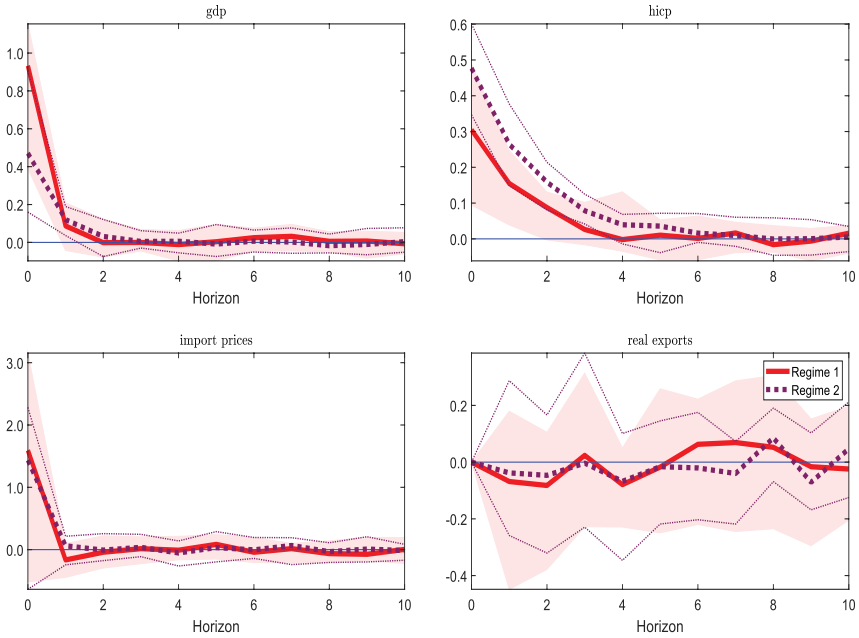
Following Chen and Lee (1995), the conditional posterior of the delay parameter  $d$  follows a multinomial distribution:

$$P(d) = \frac{L(Y_t | d, \Xi)}{\sum_{d=1}^4 L(Y_t | d, \Xi)},$$

where  $L(\cdot)$  represents the likelihood function. The model is estimated using 30,000 iterations of the Gibbs sampler, with a burn-in phase of 25,000 iterations.

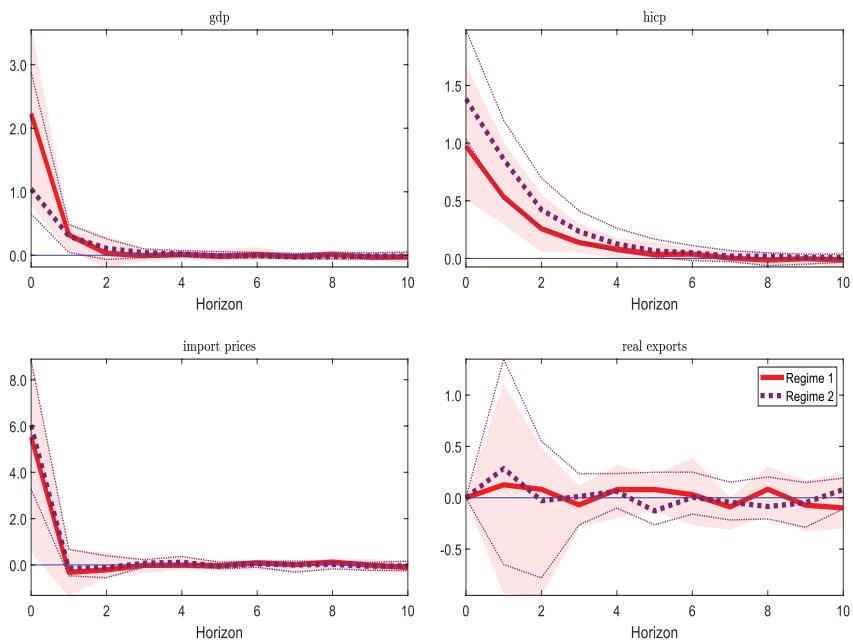
# Appendix 3: Non-Normalised IRFs from the Threshold VAR

Figure App 1: *Impulse responses to a demand shock of one standard deviation in each regime*  
(log differences)



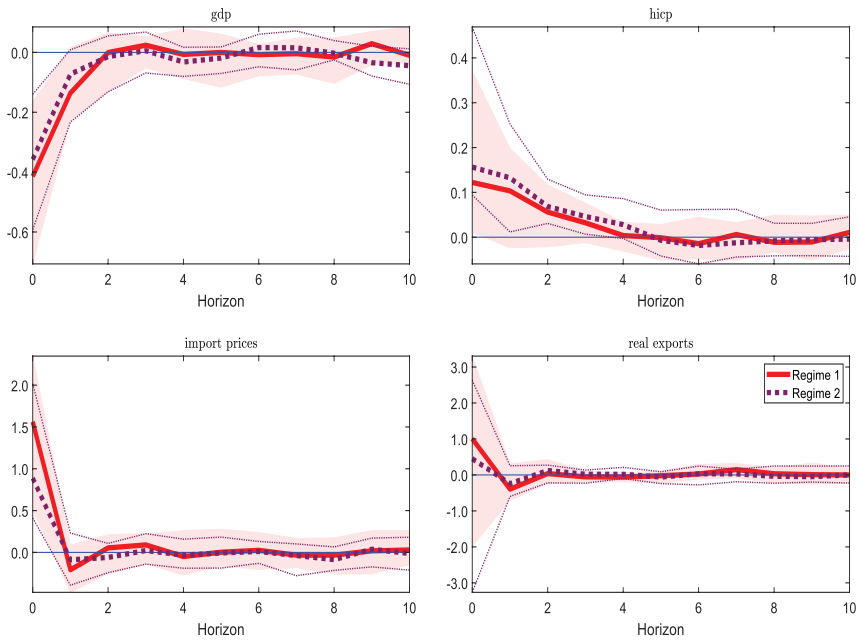
Note: The x-axis refers to the horizon of the impulse responses in quarters.

Figure App 2: *Impulse responses to a demand shock of three standard deviations in each regime (log differences)*



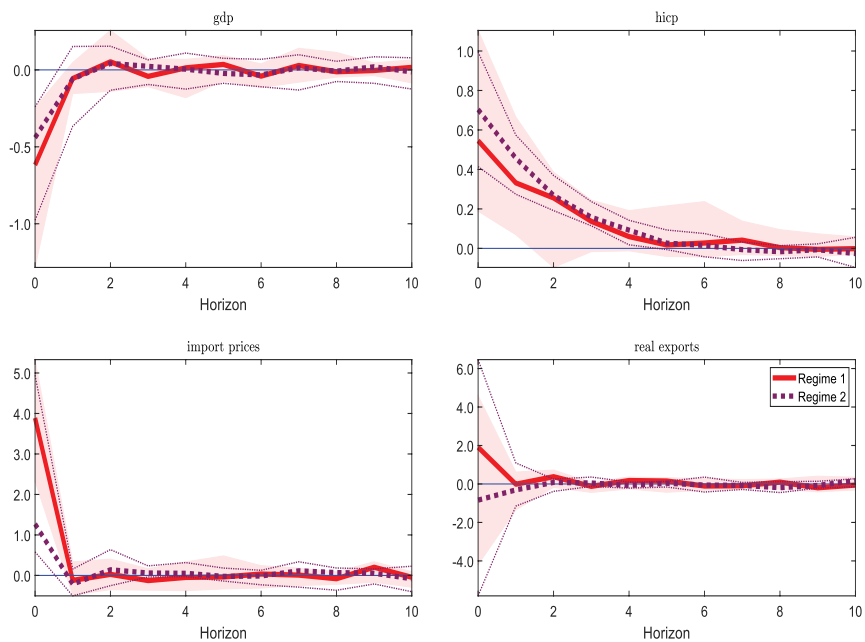
Note: The x-axis refers to the horizon of the impulse responses in quarters.

Figure App 3: *Impulse responses to a global supply (foreign price) shock of one standard deviation in each regime (normalised to a one percent increase in import prices on impact in each regime)*  
 (log differences)



Note: The x-axis refers to the horizon of the impulse responses in quarters.

Figure App 4: *Impulse responses to a global supply (foreign price) shock of two standard deviations in each regime (normalised to a one percent increase in import prices on impact in each regime)* (log differences)



Note: The x-axis refers to the horizon of the impulse responses in quarters.

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