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Food Price Inflation in Poland: Measuring Macroeconomic Shocks – A Bayesian SVAR Approach[1]

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Abstract

In the context of global warming and recurring economic crises, food prices have become an increasing challenge for households worldwide. Concerns about rising food prices, food security, and living standards are being exacerbated by the price volatility and increasing dynamics of agricultural products. Despite the importance of this issue, relatively few studies focus on European countries. The aim of this paper is to estimate the impact of global macroeconomic shocks on food prices in Poland using quarterly data for the period from Q1 2000 to Q2 2025. To this end, a Bayesian structural vector error correction model identified through sign and zero restrictions is employed to assess how such shocks have influenced the dynamics of food prices. The study highlights the dominant role of domestic shocks in explaining FCPI variability, although outside the 2019–2024 period, domestic policy and cost factors generally contributed to moderating food price inflation. The findings have important implications for policymakers and contribute to the theoretical understanding of food price dynamics.

Keywords: food price inflation, domestic shocks, external shocks, BSVAR.

1. Introduction

Poland is considered one of the European Union's largest food producers. From 2004 to 2018, Poland ranked sixth in the European Union in terms of food-industry trade value, accounting for roughly 9% of total food production [1]. At the same time, Poland's accession to the European Union has resulted in a significant increase in its links with international markets, as well as a rapid increase in food exports, transforming the country into a sustainable net food exporter. This is also reflected in Poland's growing focus on meat, cereals, fruits, and vegetables [2–4]. Food prices play a crucial role in the consumer basket in Poland. Since the early 2000s,

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their weight has slightly decreased. Nevertheless, in 2025 they still account for approximately one quarter of all prices included in the CPI, making them the most influential category.

This research makes multiple contributions to the body of literature. First, our research advances our understanding of Poland's food price inflation. The factors influencing food price inflation in Poland are thoroughly examined in this article. Poland's entry into the European Union, the global financial crisis, the global commodity crises, the food crisis of 2007–2008, a period of low inflation/deflation, the COVID-19 pandemic, the war in Ukraine, and the energy crisis are just a few of the major events covered by our study. Second, although our study is limited to Poland, it can be applied to other Central and Eastern European nations because of their similar political, historical, and economic backgrounds. Third, our work is a significant step toward future research and policy concerning the socioeconomic effects of macroeconomic shocks.

This research attempts to estimate the effects of various macroeconomic shocks, both supply-side and demand-side, on the inflation of food prices in Poland between Q1 2000 and Q2 2025. the cointegrated Bayesian structural vector autoregression model (BSVAR) in its error correction form is employed in this study. The structural shocks are identified through sign and zero restrictions and covers both domestic and external shocks, as well as demand-side and supply-side.

The article's remaining sections are arranged as follows. The literature review is presented in Section 2. The research methodology and data are presented in Section 3. The key findings from the empirical investigation are presented and discussed in Section 4. A summary and recommendations for additional research are included in Section 5's conclusion. Finally, Appendices A–C contain supplementary material.

2. Literature review

In recent years, the issue of food security has gained increasing attention. While earlier research largely focused on developing regions, such as Sub-Saharan Africa, India, or Pakistan, a combination of economic shocks and climate change has raised similar concerns in European Union [5]. These developments have motivated a growing body of work examining food security. For example, the food security index proposed by Matejková et al. [6] shows that rising food prices constitute the primary factor undermining food security across Europe. The most vulnerable countries in this regard are those located in Central and Eastern Europe, including Poland.

Food prices can vary depending on a variety of factors. The supply-demand ratio continues to be the main factor influencing food prices [7, 8]. The degree of socioeconomic development of a nation also has a big impact on food prices because rising social affluence causes changes in consumption habits. Additionally, government subsidies, fertilizer prices, rise in commodity prices, energy costs, and the effectiveness of production systems all have an impact on food prices [9, 10].

Food prices on global markets increased slowly between 2000 and 2002 as a result of the economic downturn and declining fuel prices. When the cost of energy resources, especially crude oil, skyrocketed in 2003, this was altered. This resulted from the Middle East's tense geopolitical situation. In addition, the drought that Europe was experiencing at the time caused agricultural output to drop. Since 2004 saw the biggest expansion in the history of the European Union, it was a watershed year for Poland and the entire Central and Eastern European region. However, the European Union's expansion to include ten new member states was a factor in the increase in food prices worldwide. Additionally, the state of the world economy improved, which raised demand for investments and consumers [11].

Prices of agricultural commodities experienced unusually high volatility between 2007 and 2009. Raw food prices increased significantly between May 2007 and February 2008 before stabilizing by March 2009. The asymmetric character of the food supply chain's price transmission mechanism was made clear during this time. Producer prices increased by 9% and consumer prices increased by 5% as a result of the food crisis of 2007–2008. However, small producers and retailers were the only ones who profited from the 2009 decline in agricultural commodity prices. As a result, retail food prices continued to be high, which hurt European consumers and reduced the competitiveness of the European food market [12].

Following the 2008-2009 financial crisis and the eurozone debt crisis, economists were left with many questions about the European Union's inflationary trajectory [13]. In addition to both crosses, the oil sector and fluctuating oil prices have contributed to relatively low inflation. Between April 2014 and January 2015, oil prices dropped by nearly 50%. Agricultural commodities fell only 7% over the same time period [13, 14]. However, the price of dairy products, especially butter, skyrocketed in the second decade of the twenty-first century. Poland is a major producer of dairy products, just like the United States, Australia, New Zealand, and the rest of European Union members. It manufactures a variety of dairy products to satisfy both domestic and international markets [15–18]. Global trade in dairy products started to expand dramatically in the mid-2000s, mainly due to shifting consumption patterns in China, the country's population's increasing affluence, and the growth of the middle class. As a result, China saw the biggest increase in dairy product imports between 2005 and 2018 [19]. Furthermore, as misconceptions about butter's detrimental effects have been dispelled, attitudes toward it have shifted in developed nations. Additionally, food produced in the European Union was subject to an embargo by Russia in 2014, which led to a decrease in production. A decrease in the supply of milk was also caused by the European Union's elimination of milk quotas. New Zealand, the biggest exporter of butter in the world at the time, was also modernizing its dairy sector [17, 20, 21]. Significant correlations between butter prices in Poland, a few EU nations, and Oceania were also confirmed by Domagała [20] research.

Global food security has been adversely affected by the pandemic. Countries with high levels of food security have been particularly affected in terms of food availability and affordability [22, 23]. The COVID-19 pandemic has created a dynamic and uncertain economic environment, which has affected the inflation of food prices. At the beginning of the pandemic, there was a significant increase in food prices (especially meat, dairy products, and canned/frozen vegetables and fruit), driven by numerous restrictions and panic among the public, who began stockpiling food while store shelves were empty [24–28]. Inadvertently, restrictions put in place to stop the pandemic's spread have made food prices rise. During the pandemic, food prices have also increased as a result of the disruption of food supply chains, forced closure of business or labour shortage [22, 24]. A shift in consumer behaviour has also been observed during the COVID-19 pandemic period. For example, Poland has seen a rise in interest in organic food as a result of health concerns and a focus on environmental issues [29, 30].

With the start of the Russian-Ukrainian war in February 2022, the COVID-19 pandemic's effects became more severe [31–33]. Poland and other nearby nations are said to have been most severely affected by the Russian-Ukrainian War, especially compared to Western EU – see e.g. Daianu et al. [34]. First, the demand for necessities rose as a result of the Ukrainian refugee crisis. Furthermore, because Russia and Belarus were major exporters of fertilizer and agricultural products, the sanctions placed on them upset the world food system. In contrast, Ukraine saw a reduction in food production, losses in food supplies and production, and a lack of workers in rural areas. Around 18% of Poland's consumer basket was made up of food expenses in July 2022, when food price inflation hit 14%. It should be noted, however, that Poland has experienced lower food price inflation than other Central and Eastern European nations, a consequence of its reduced reliance on imports as Poland is the EU's third-largest net exporter of food products [31, 35–38]. However, energy prices rose sharply as a result of the conflict in Ukraine, raising the price of producing food [39, 40].

3. Empirical strategy and data

This study aims to estimate the impact of several macroeconomic shocks, both demand and supply, on food price inflation in Poland. The data used for the analysis in this study is quarterly seasonally adjusted[3] data ranging from Q1 2000 to Q2 2025 obtained from several sources (Table 1. briefly describes all the variables and sources used for the analysis and Figure 1. provides time series plot of the variables – see Appendix A.). The data set was selected based on literature focused on modelling inflation (and food price inflation specifically), where both intra-country and global variables were used – see e.g. Akinbode et al. [42], Anderl & Caporale [43], Ascari et al. [44], Samal et al. [10].

Our data set includes domestic food price inflation (*FCPI*), domestic core inflation (*CCPI*), real GDP *per capita* (*RGDP*), broad money (*M3*), real effective exchange rate (*REER*), and the global-side variables – real food prices (*FFPI*), oil prices (*OP*) and fertilizer prices (*FP*). To gain deeper knowledge about the reaction of food price inflation in Poland to the selected shocks, we use the structural vector error correction (SVEC) model – see Appendix A. To examine the impact of the macroeconomic shocks on food price in Poland we restore to impulse response functions (IRF), variance decomposition of forecast errors (FEVD) and historical decomposition (HD) methods.

We use both sign and zero restrictions to fully identify six shocks. On the demand side, we distinguish between fiscal policy and monetary policy. On the supply side, we differentiate between domestic cost-push shock, global food price shock, oil price shock and fertilizer price shock. Therefore, this exercise delivers novel estimates of the food price inflation response to demand and supply shocks and informs about the relative importance of each of these shocks in driving the recent burst in Polish food price inflation.

Our identification strategy is discussed below and outlined in Table 2. The sign restrictions should be seen as in line with economic theory. Our assumptions are summarized as follows. (1) A positive fiscal shock boosts output and subsequently increases food price inflation and core inflation; see e.g. Abbas & Waheed [49], Carrasco & Mukhopadhyay [50], Samal et al. [10], Sasmal [51]. We also assume it will boost the REER due to higher overall inflation. Also, domestic shocks should not affect global side variables. (2) A positive monetary shock boosts money supply, output, food price inflation, core inflation and lowers the REER; see e.g. Anderl & Caporale [43]. (3) Adverse domestic cost-push shocks reduce output and raise food price inflation and core inflation; see e.g. Anderl & Caporale [43]. (4) A positive food price shock boosts both domestic and global food prices and reduces output. We assume it does not affect oil or fertilizer prices. (5) A negative oil supply shock boosts both domestic and global food prices. Also, we assume it results in a surge of fertilizer prices; see e.g. Sanyal et al. [52]. We also assume it should raise a core inflation; see e.g. Bańbura et al. [53], Conflitti & Luciani [54]. (6) A negative fertilizer supply shock increases both domestic and global food prices and does not affect oil prices.

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[3] The data were adjusted using seasonal R-package [41].

Table 2. Restrictions for Identifying Shocks

	Fiscal policy	Monetary policy	Cost-push	Global food prices	Oil prices	Global fertilizer prices
FCPI	+	+	+	+	+	+
CCPI	+	+	+		+	
FFPI	0	0	0	+	+	+
RGDP	+	+	-	-	-	-
M3		+				
REER	+	-			-	
OP	0	0	0	0	+	0
FP	0	0	0	0	+	+

Note: An entry with +/-/0 denotes a positive/negative/zero response of the variable (rows) to the specific structural shock (columns). An empty cell implies an unrestricted response.

4. Results and discussion

This section presents the main results and it is arranged as follows: firstly, we present the chosen set of hyperparameters of prior distribution; secondly, we discuss the results of model based on the IRF's; thirdly, we elaborate on the relative importance of shocks using FEVD's and HD's.

We set the following values of the hyperparameters of the prior distribution:

- $\sigma_i \sim iG(1{,}0.01)$, for $i = 1,2,\dots,n$,
- $$\lambda_{ij}|h_{ij} \sim \begin{cases} N(0,h_{ij})\mathbf{1}_{[\lambda_{ij}>0]}, & \text{if } M_{ij} = 1 \\ N(0,h_{ij})\mathbf{1}_{[\lambda_{ij}<0]}, & \text{if } M_{ij} = -1 \\ \delta_0(\lambda_{ij}), & \text{if } M_{ij} = 0 \\ N(0{,}0.1), & \text{otherwise} \end{cases},$$

with $$M = \begin{bmatrix} 1 & 1 & 1 & 1 & 1 & 1 \\ 1 & 1 & 1 & & 1 & \\ 0 & 0 & 0 & 1 & 1 & 1 \\ 1 & 1 & -1 & -1 & -1 & -1 \\ & 1 & & & & \\ 1 & -1 & & & -1 & \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 & 1 \end{bmatrix}$$ and $h_{ij} \sim iG(3{,}100)$,

for $i = 1,2,\dots,n,\ j = 1,2,\dots,s$,

- $\Phi|\Sigma \sim mN(0,\Sigma,0.1I_{r+n(p-1)+l_D})$,
- $\beta^* \sim mN(0,I_r,\frac{1}{10m}I_m)$.

All results presented below are based on 50 000 MCMC draws, preceded by as many *burn-in* iterations.

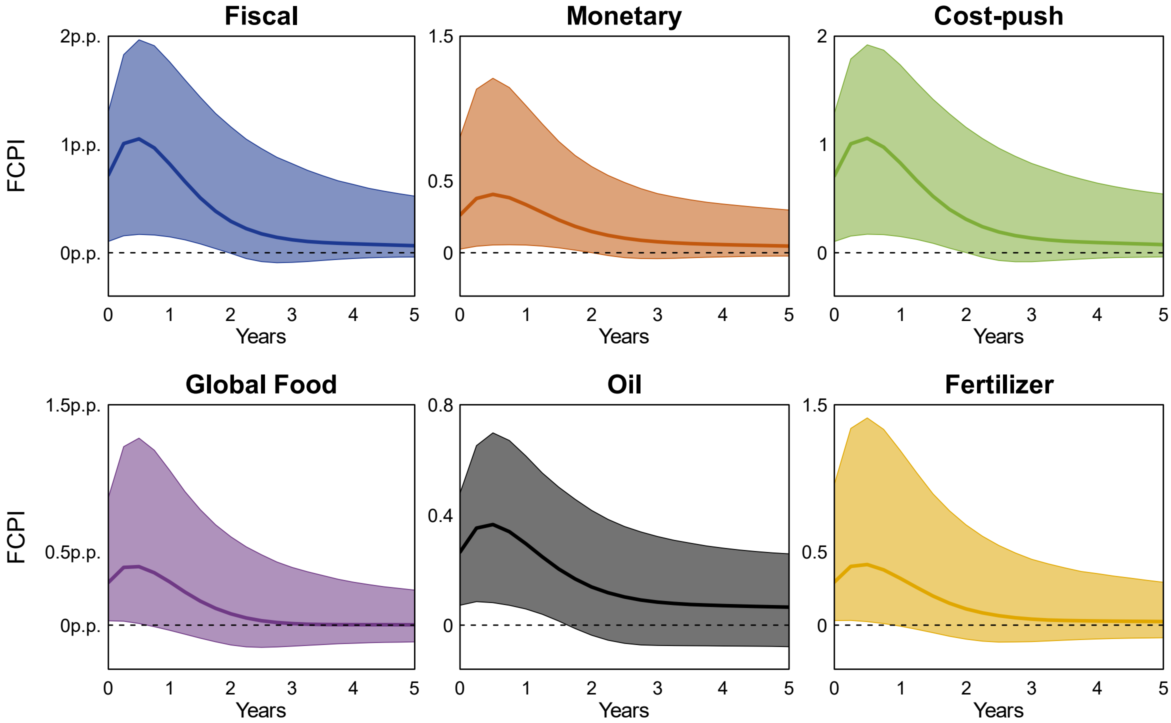
Regarding the cointegrating relationships, we relied on the Johansen [55] procedure rather than on formal Bayesian approaches based on model comparison. The results of this stage of the analysis indicated the presence of three cointegrating vectors (r = 3), and this number of relationships was therefore adopted for the subsequent analysis.

Figure 2. presents the IRFs of food price inflation – as it is the variable of interest – to all the shocks that we identify in the model. However, we will discuss the responses of all variables whose IRFs are available in Appendix B (see Figure 3.).

Given the imposed sign restrictions, a fiscal policy shock exerts a positive effect on food price inflation, core inflation, real GDP per capita, and the REER. The shock induces an immediate increase in food price inflation, with the response remaining statistically significant for the subsequent 2 years. A similar pattern is observed for CCPI. By contrast, the shock has a persistent effect on the REER, whereas the response of real GDP is confined to the impact period. The response of M3 remains statistically insignificant. The remaining variables, consistent with the identifying assumptions embedded in the contemporaneous impact matrix, do not display statistically significant reactions to the fiscal shock.

A comparable pattern emerges following a monetary policy shock. The shock generates an immediate positive effect on food price inflation, which remains statistically significant for 2 years. CCPI also responds positively, with the effect persisting for an additional 3 years. The disturbance produces only an impact increase in real GDP and M3, alongside a decline in the REER. The other variables do not exhibit statistically significant responses to the shock, neither at impact as restricted by identification, nor over the longer horizon.

Figure 2. Impulse Response Functions of Food Price Inflation

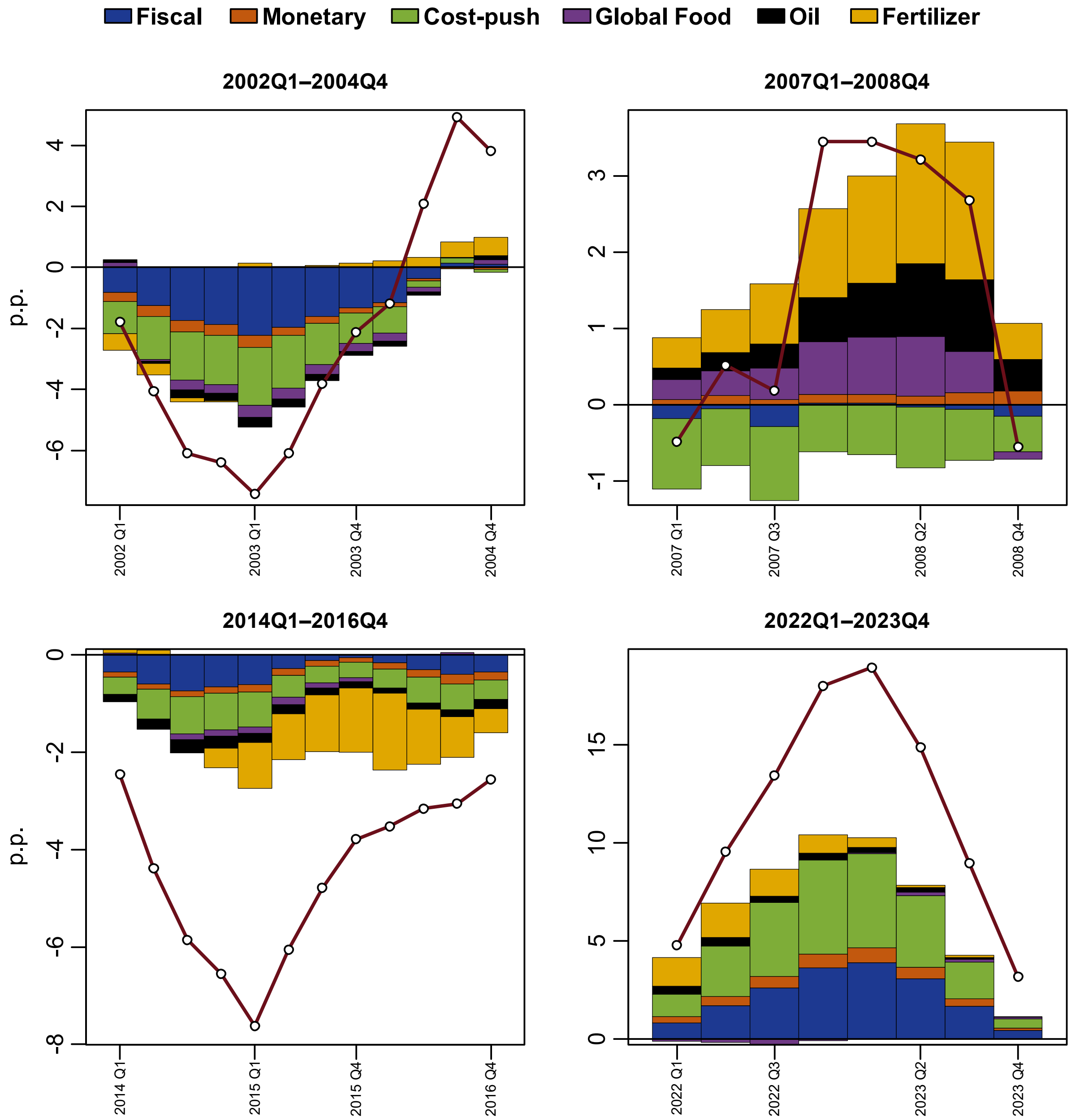


Source: Own elaboration; access to data 18 September 2025.
Note: Impulse responses to one-standard-deviation shocks. Solid lines represent the posterior pointwise median, and the shaded areas indicate the 5th and 95th percentiles of the impulse responses.

Both shocks primarily pertain to the domestic demand side and appear to reflect variable dynamics consistent with the empirical literature examining the effects of fiscal and monetary policy on food price inflation and inflation more broadly; see e.g. Adjemian et al. [7], Anderl & Caporale [43], Ascari et al. [44], Samal & Goyari [56]. However, it is good to mention that e.g. Bareith & Fertő [57] do not provide results confirming statistically significant results of monetary policy shocks. Some analysis suggests even that the monetary tightening (i.e. negative monetary shock in our convention) can led to rise of food price inflation due to e.g. higher cost of capital – see e.g. Bhattacharya & Jain [58]. There is also a gap in literature regarding response of food price inflation to fiscal shocks.

The cost-push shock generates an immediate increase in food price inflation, as implemented via restrictions, with the response remaining statistically significant for approximately 2 years. Core inflation rises for roughly 2.5 years. The response of real GDP is persistent and negative for 2 years, while the remaining variables do not exhibit statistically significant reactions. Notably, despite allowing for unrestricted responses of the money supply and exchange rate, neither variable reacts significantly to the cost-push shock. These findings are in line with the results reported by Anderl and Caporale [43].

Figure 5. Historical Decompositions of the Deviation from Average Food Price Inflation



Source: Own elaboration; access to data 18 September 2025.
Note: The maroon line shows the deviation of food price inflation from its average. Only structural shocks (the pointwise median) are included, while idiosyncratic shocks are excluded; as a result, the cumulative contributions do not always sum to the total FCPI.

A global food price shock generates an immediate increase in food price inflation in Poland; however, its persistence is considerably shorter, dissipating after approximately two quarters. These findings are broadly consistent with previous studies examining the transmission of global food prices to domestic price dynamics; see e.g. Anderl & Caporale [43], Davidson et al. [59], El Ghin & El-Karimi [60], Peersman [61], Samal et al. [10]. The shock raises the level of global food prices for two quarters and lowers the RGDP for about a year. After approximately 1.5 years shock significantly raises the REER. Other variables do not response significantly to the shock.

Food price inflation in Poland responds positively to an oil price shock for more than 1.5 years following the disturbance. These results are consistent with part of the existing literature; see e.g. Balogh & Sárvári [62], Mishra et al. [63], whereas Anderl and Caporale [43] and Baumeister and Kilian [64] report no statistically significant response of this variable. In contrast, the decline in RGDP is confined to the impact period and is not sustained over time.

Finally, a shock originating from global fertilizer prices translates into an immediate increase in food price inflation, with a persistence comparable to that observed in the case of a global food price shock. There is a significant gap of studies covering linkage between fertilizer prices and the food price inflation in broad sense. However, there are some studies examining reaction of single-type product prices and supply; see e.g. Brunelle et al. [65], Mishra et al. [66]. The disturbance also induces a rise in global food prices for about 2 quarters, while exerting a small negative effect on RGDP for about a year.

The FEVD plots (see Appendix C – Figure 4.) provide additional insights derived from the model. It can be observed that domestic shocks, in particular fiscal and cost-push shocks, dominate other factors in explaining the forecast error variance of food price inflation. A similar pattern is evident for CCPI. The remaining variables primarily respond to global disturbances, especially oil price shocks.

At the same time, based on the historical decompositions for selected periods (see Figure 5.), domestic shocks appear to be the main drivers of food price inflation dynamics. However, outside the 2019–2024 period (see Figure 6. in Appendix D), fiscal and monetary policy, as well as domestic cost factors, exerted downward pressure on food price inflation.

5. Concluding remarks

The aim of this study was to estimate the impact of demand and supply-side shocks on food inflation in Poland using quarterly data for the period from Q1 2000 to Q2 2025. We identified six shocks using sign and zero restrictions and applied BSVEC model, alongside with IRFs, FEVDs and historical decomposition analysis to examine linkage between the shocks and the food price inflation.

Overall, food price inflation in Poland is driven primarily by domestic shocks, particularly fiscal and cost-push disturbances. Both fiscal and monetary shocks generate an statistically significant increase in FCPI, with effects persisting for around two years. Cost-push shocks likewise produce a sustained rise in FCPI of comparable duration. Among external factors, oil price shocks exert the most persistent upward pressure on FCPI, whereas global food and fertilizer price shocks have shorter-lived effects that dissipate relatively quickly. Forecast error variance and historical decompositions further confirm the dominant role of domestic shocks in explaining FCPI variability, although outside the 2019-2024 period domestic policy and cost factors generally contributed to moderating food price inflation.

A limitation of the present study is that it provides evidence for a single country only. Extending the analysis to a broader set of economies – particularly within the CEE region – could yield valuable insights for the design and coordination of regional policy responses. Another promising avenue for future research would involve disaggregating the food price inflation index into specific product categories, allowing for a more granular assessment of transmission mechanisms. Additionally, the application of VAR-GARCH-type specifications would enable modelling of time-varying volatility, or heteroskedasticity of errors and could facilitate the examination of volatility spillovers and contagion effects between food commodity markets and energy or fertilizer markets.

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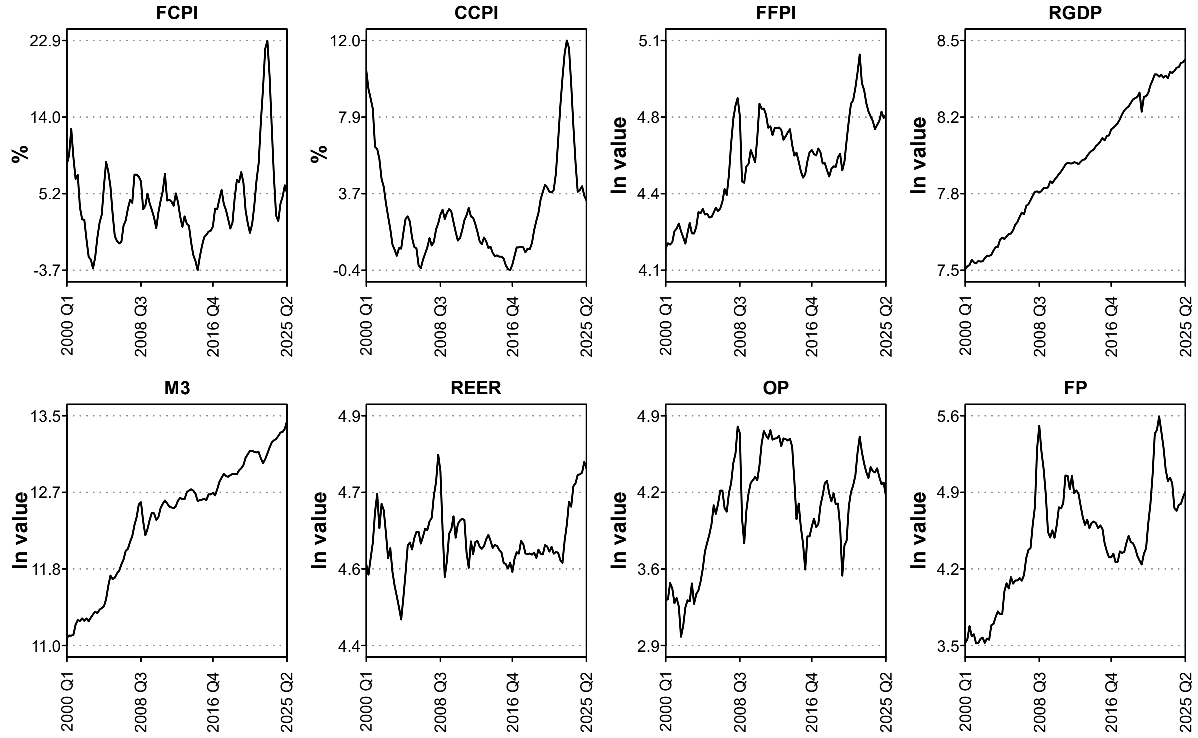
Appendix A – Data and Methods Description

Table 1. Variable Description

Variable	Description	Source
FCPI	Year-on-year inflation based on CPI index for the food and non-alcoholic beverages basket.	OECD
CCPI	Year-on-year core inflation based on CPI index, all items non-food non-energy.	OECD
FFPI	Global real price index for 5 food product groups. Averaged prices from 2014-2016 = 100. Adjusted seasonally, logarithmic values used.	FAO
RGDP	GDP *per capita* at constant 2015 prices, originally in euro. Converted to USD using the average EUR/USD exchange rate for 2015 based on ECB data. Adjusted seasonally, logarithmic values used.	Eurostat
M3	Money supply (M3 aggregate), originally in national currency. Converted to USD using the average exchange rate at each quarter based on FRED data. Adjusted seasonally, logarithmic values used.	OECD
REER	Real effective exchange rate; aggregated from the exchange rates for 42 most important trading partners. Averaged exchange rates from 2015 = 100. Increase in value - real appreciation, deterioration in competitiveness. Adjusted seasonally, logarithmic values used.	Eurostat
OP	Nominal Brent crude oil prices – USD per barrel, logarithmic values used.	FRED
FP	Index of nominal prices of four ingredients used in fertilizers: Natural Phosphate Rock; Phosphate; Potassium; Nitogenous, created in the ratio: 16.9/21.7/20.1/41.3. Average prices from 2010 = 100, logarithmic values used.	World Bank

Source: Own elaboration; access to data 18 September 2025.

Figure 1. Time Series Plots



Source: Own elaboration; access to data 18 September 2025.

We start with an n-variate vector autoregressive process $y_t = (y_{t1}, \dots, y_{tn})'$ of order p ($y_t \sim VAR(p)$), in the structural vector error correction form, with deterministic terms:

$$\Delta y_t = \alpha\beta'\tilde{y}_{t-1} + \sum_{i=1}^{p-1}\Gamma_i\Delta y_{t-i} + \Psi D_t + \varepsilon_t, \quad \varepsilon_t \sim iiN(0,\Omega), \tag{1}$$

where $\tilde{y}'_{t-1} = (y'_{t-1} \quad d'_{t-1})$, d_{t-1} denotes deterministic components restricted to cointegrating relations, α in an $n \times r$ matrix of adjustment coefficients, $\beta' = (\tilde{\beta}' \quad \phi')$ is an $r \times m$ cointegration matrix ($m \geq n$) of a full column rank r, r is the number of cointegration relationships ($0 < r < n$, if they exist), or $r = 0$ (for a non-stationary but non-cointegrated system), or $r = n$ (for a stationary VAR), the l_D- dimensional vector D_t consists of unrestricted deterministic variables (such as constant, deterministic trends, seasonal dummies; see, e.g., Juselius [45], and Ω is a covariance matrix of the reduced form shocks (ε_t). Note that for any non-singular matrix $M_{r\times r}$ we can obtain the equivalent representation of the matrix $\Pi = \alpha\beta'$, i.e. $\alpha\beta' \equiv \alpha^*(\beta^*)'$, where $\alpha^* = \alpha M,\ \beta^* = \beta(M')^{-1}$. We take this indeterminacy into account by assuming that β has orthonormal columns, so it is an element of the space of orthonormal r-frames in $\mathbb{R}^m$, which is called the Sitefel manifold ($V_{r,m}$). This assumption allows us to identify the cointegrating space.

To simplify the notation we rewrite the model (1) in the following form

$$\Delta y_t = \Phi x_t + \varepsilon_t, \quad \varepsilon_t \sim iiN(0,\Omega), \tag{2}$$

where $\Phi = (\alpha \quad \Gamma_1 \quad \dots \quad \Gamma_{p-1} \quad \Psi)$ and $x'_t = (\tilde{y}'_{t-1}\beta \quad \Delta y'_{t-1} \quad \dots \quad \Delta y'_{p-1} \quad \mathrm{D}'_{\mathrm{t}})$.

We use the idea of Korobilis [46] and assume that the reduced-form shocks are driven by a few common orthonormal factors, and we impose identification restrictions on their loadings, so we decompose the reduced-form errors (ε_t) using the following factor model specification

$$\varepsilon_t = \Lambda f_t + \nu_t, \tag{3}$$

where Λ is an $n \times s$ matrix of factor loadings, f_t is an $s \times 1$ vector of orthonormal factors, i.e. $f_t' f_t = I_s$, and ν_t is an $n \times 1$ vector of idiosyncratic shocks. In summary, we decompose n reduced-form shocks into $s + n$ shocks, of which the s common shocks (f_t) are considered structural and the ν_t comprises n nuisance shocks. The factors and idiosyncratic shocks are independent contemporaneously and, in all leads, and lags ($f_t \perp \nu_\tau,\ t \neq \tau,\ t,\tau \in \{1,2,\dots,T\}$). We assume that $\nu_t \sim iiN(0,\Sigma)$, where Σ is an $n \times n$ diagonal matrix, and $F_{T\times s} = (f_1 \quad f_2 \quad \dots \quad f_T)'$ is the matrix of structural shocks uniformly distributed on the space of orthonormal s-frames in $\mathbb{R}^T$, i.e. the Stiefel manifold $V_{s,T}$. We can decompose the covariance matrix of the reduced-form shocks as $\Omega = \Lambda\Lambda' + \Sigma$. To fully identify the structural model we impose the restrictions on Λ.

We impose priors on the parameters of the following representation of our model

$$\Delta y_t = \Phi^* x^*_t + \Lambda f_t + \nu_t, \tag{4}$$

where $\Phi^* = (\alpha^* \quad \Gamma_1 \quad \dots \quad \Gamma_{p-1} \quad \Psi)$, $x^*_t = (\tilde{y}'_{t-1}\beta^* \quad \Delta y'_{t-1} \quad \dots \quad \Delta y'_{p-1} \quad \mathrm{D}'_{\mathrm{t}})'$, α^* is the non-normalized matrix of adjustment coefficients, β^* is the nonnormalized matrix of cointegrating vectorsWe assume that the normalized cointegrating vectors are the orientation part of a polar decompositions of β^*, so $\beta = \beta^*\left(\beta^{*\prime}\beta^*\right)^{-1/2}$ and $\alpha = \alpha^*\left(\beta^{*\prime}\beta^*\right)^{1/2}$.

The following list provides details about the prior densities imposed on the model's parameters:

- inverse gamma distributions for the variances of idiosyncratic shocks: $\sigma_i \sim iG(\rho_i, \kappa_i)$[4],
- a (truncated-) normal for nonzero factor loadings:

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[4] By $iG(a,b)$ we denote an inverse gamma distribution with mean $\frac{b}{a-1}$ (under $a > 1$), and variance $\frac{b^2}{(a-1)^2(a-2)}$ (under $a > 2$).

$$\lambda_{ij}|h_{ij} \sim \begin{cases} N(0, h_{ij})\mathbf{1}_{[\lambda_{ij}>0]}, & \text{if } M_{ij} = 1, \\ N(0, h_{ij})\mathbf{1}_{[\lambda_{ij}<0]}, & \text{if } M_{ij} = -1, \\ \delta_0(\lambda_{ij}), & \text{if } M_{ij} = 0, \\ N(0, h_{ij}), & \text{otherwise,} \end{cases}$$

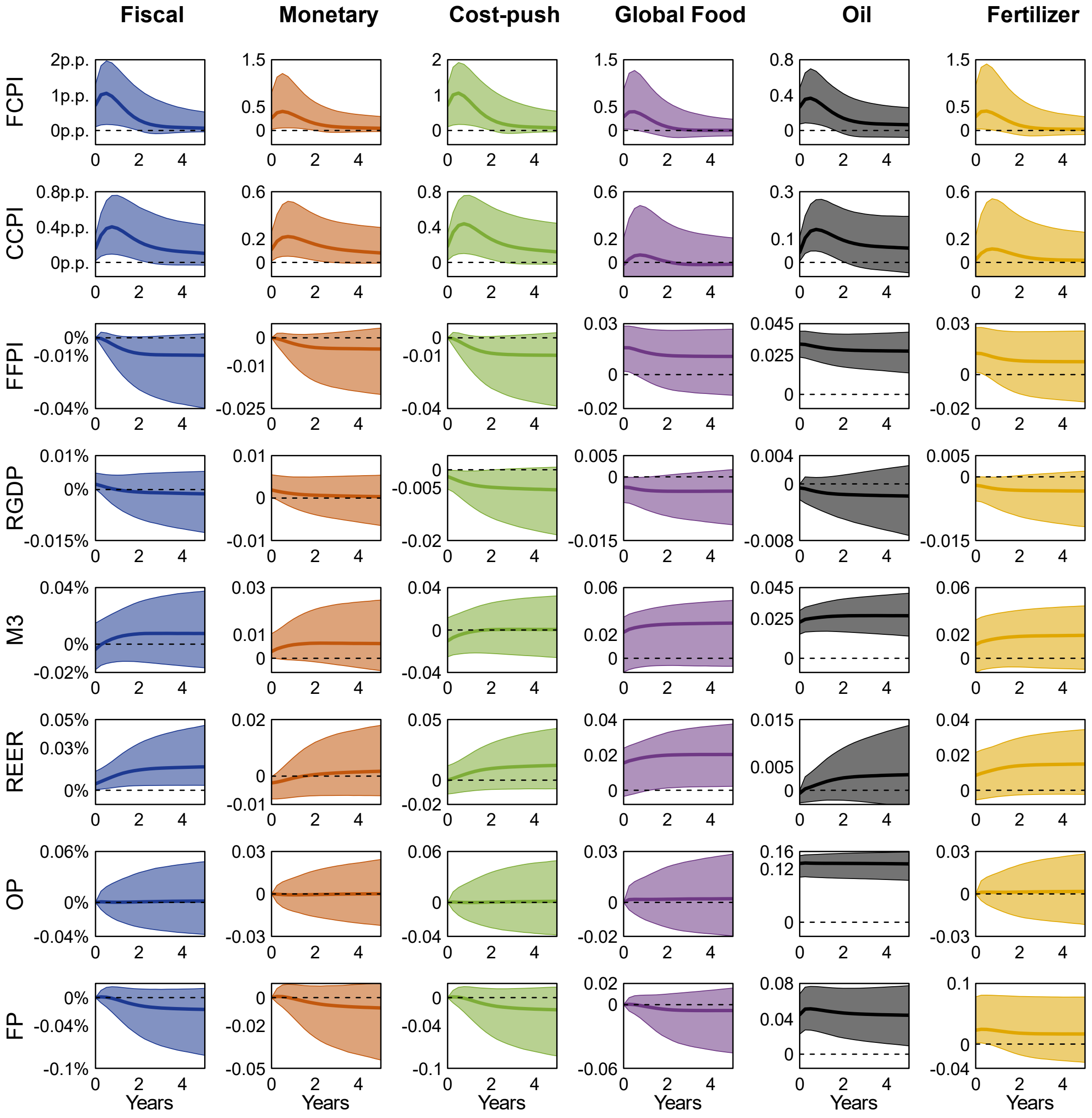
for $i = 1,2,\dots,n,\ j = 1,2,\dots,s$, $\delta_0(\lambda_{ij})$ is the Dirac delta function for λ_{ij} at zero (i.e. a point mass function with all mass concentrated at zero) and $\mathbf{1}_{[\cdot]}$ is an indictor function, all sign and zero restrictions imposed on Λ are collected in the matrix M, with entries +1 for positive signs, −1 for negative signs, 0 for zero restrictions, and a missing value for no restriction; the hyperparameters may be deterministically set by the researcher or modelled with an inverse gamma hyperprior distributions, $iG(n_{ij}, s_{ij})$.

- a matrix normal distribution for the matrix Φ^*: $\Phi|\Sigma \sim mN(0, \Sigma, \underline{\Omega}_{\Phi})$, leading to $\phi^*|\Sigma \sim N(0, \Sigma \otimes \underline{\Omega}_{\Phi})$, where $\phi^* = vec(\Phi^*)$,
- a matrix normal distribution for the matrix of non-normalized cointegrating vectors: $\beta^* \sim mN(0, I_r, P)$, leading to $\beta \sim MACG(P)$ (a matrix angular central Gaussian distribution with an $m \times m$ matrix hyperparameter P; see Chikuse [47, 48], and to $b^* = vec(\beta^*) \sim N\,(0, I_r \otimes P)$, the prior information about the cointegration space may be incorporated *via* the matrix P. In the case with no prior information $P = I_m$. Note that the $MACG$ distribution is invariant with respect to multiplication by a positive constant, i.e. $MACG(P) = MACG(cP)$, where $c > 0$.

The joint priori distribution of Φ^* and β^* is truncated by the restriction on the spectral radius $\rho(\boldsymbol{A})$ of the companion matrix $\boldsymbol{A} = \boldsymbol{A}(\Phi^*, \beta^*)$ to be less or equal one.

Appendix B – Impulse Response Functions

Figure 3. Impulse Response Functions

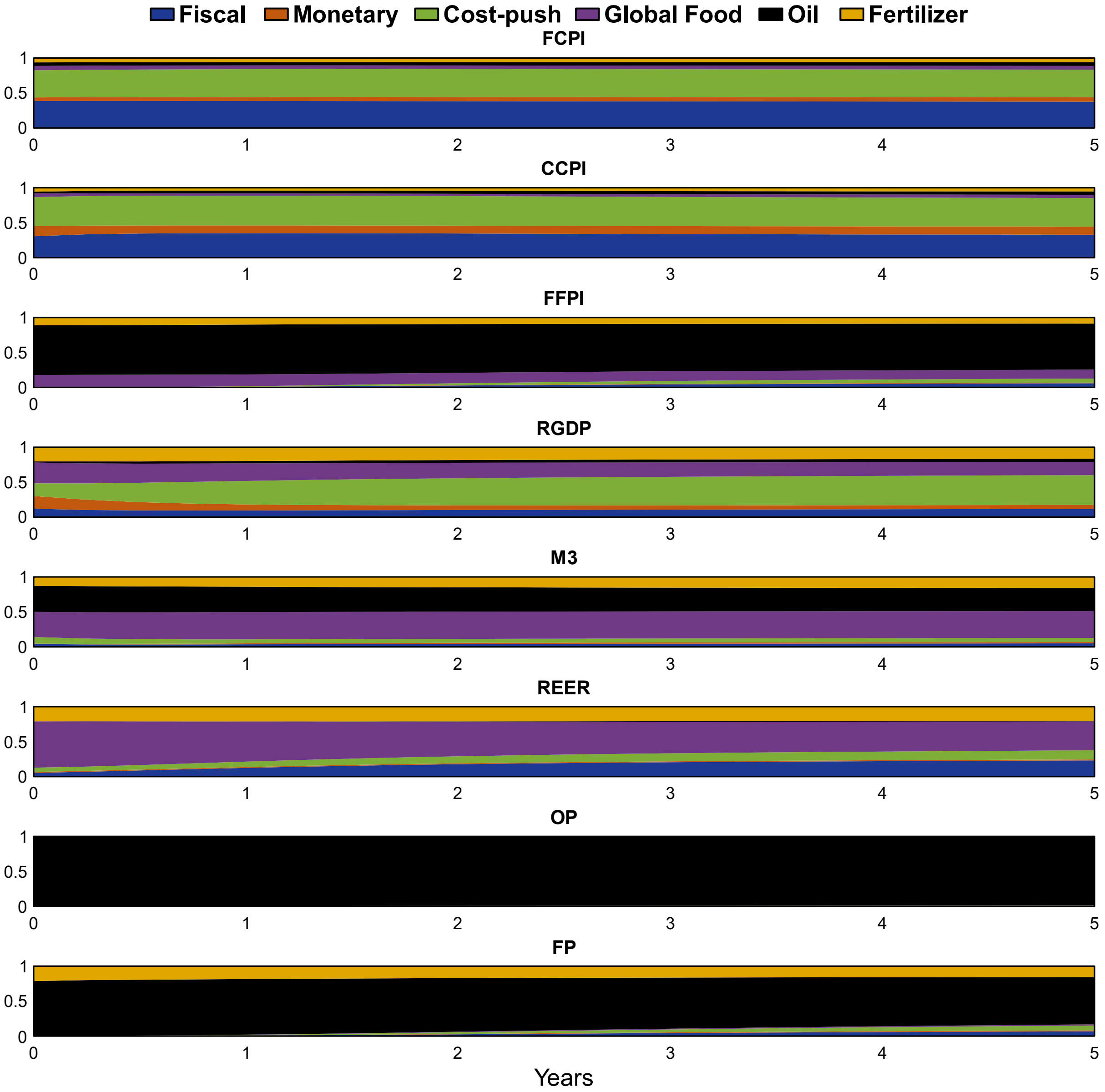


Source: Own elaboration; access to data 18 September 2025.

Note: Impulse responses to one-standard-deviation shocks. Solid lines represent the posterior pointwise median, and the shaded areas indicate the 5h and 95th percentiles of the impulse responses.

Appendix C – FEVD Plots

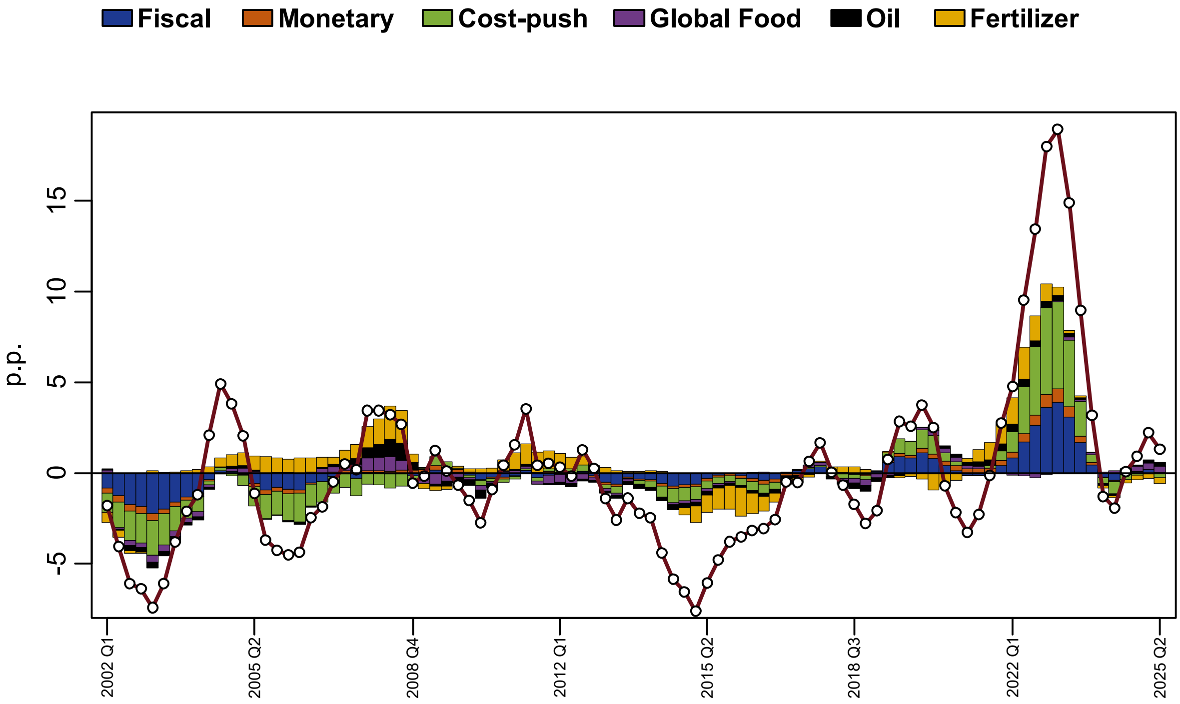
Figure 3. Variance Decomposition of Forecast Errors



Source: Own elaboration; access to data 18 September 2025.
Note: The pointwise median contributions are reported. They are normalized so that they add up to 100.

Appendix D – Historical Decompositions

Figure 5. Historical Decompositions of the Deviation from Average Food Price Inflation – Full Sample



Source: Own elaboration; access to data 18 September 2025.
Note: The maroon line shows the deviation of food price inflation from its average. Only structural shocks (the pointwise median) are included, while idiosyncratic shocks are excluded; as a result, the cumulative contributions do not always sum to the total FCPI.