Monetary Policy and Food Inflation in Central Europe: Evidence from the Visegrad Countries

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Abstract

This study examines the relationship between monetary policy and food inflation in the Visegrad Group, using monthly data and applying both OLS and quantile regression methods. Because the model is estimated in first differences and includes a three-month lag of the policy rate, all results reflect short-run month-to-month dynamics of food inflation. The analysis reveals that the monetary policy rate is significantly associated with food inflation across several quantiles, with stronger effects observed during periods of higher inflation. The study also examines the roles of exchange rates, industrial and transport inflation, with a robustness check replacing transport inflation with energy prices. This adjustment confirmed the relevance of energy prices in food inflation dynamics. The results indicate that while monetary policy does affect food prices, its effectiveness depends on the level of inflation and underlying supply-side factors. Quantile regression proves to be a valuable tool in capturing these heterogeneities. These findings can support policymakers in designing more responsive and effective strategies to manage food inflation under varying economic conditions.

Keywords

Visegrad group countries, food inflation, monetary policy, quantile regression, OLS.

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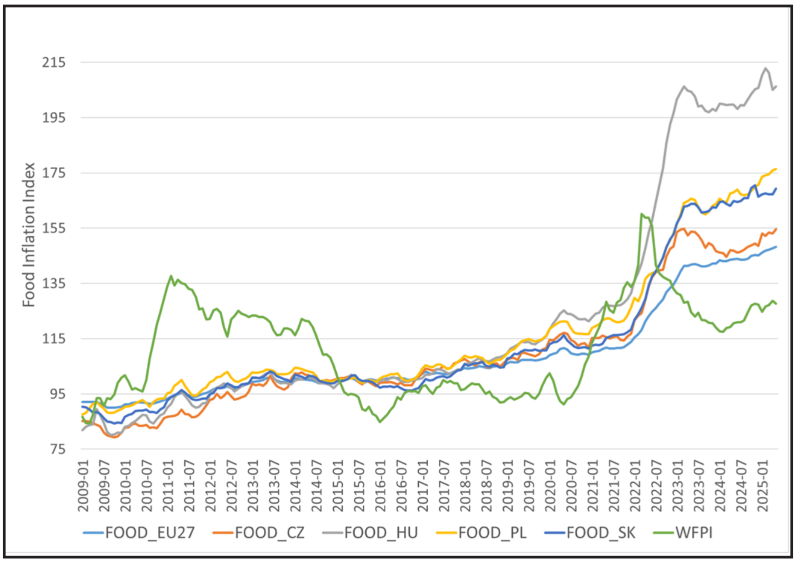
Introduction

Inflation represents one of the most important macroeconomic indicators (Mügge, 2016). Its central role is further reflected in the strategic frameworks of many central banks, which identify price stability as the primary objective of their monetary policy. From a policy perspective, food inflation poses a unique and growing challenge, particularly because of the monetary policy instruments, and especially interest rates, which have shown a measurable effect in influencing food inflation dynamics. For instance, Bhattacharyya and Jain (2020) show that an unexpected monetary tightening tends to have a statistically significant and positive impact on food prices, both in advanced and emerging markets. Also, Kaur (2023) or Sami and Makun (2024) present how monetary policy responses to food inflation vary across countries and reviews growing academic interest in this area. Moreover, De Gregorio (2012) highlights that food inflation, more than energy inflation, tends to exert stronger propagation effects on core inflation, underscoring its importance for price stability. Despite growing research on the relationship, a recent study by Fertő and Bareith (2024) emphasizes that food inflation continues to receive relatively limited attention in the literature.

Why is the issue of food inflation particularly important in the countries of the Visegrad Group (V4)? Despite their geographic and economic proximity, the Visegrad countries operate under different monetary policy frameworks. From the monetary policy standpoint, Slovakia has been regulated by the European Central Bank since 2009, whereas Czechia, Hungary, and Poland retain independent central banks with their own inflation targeting strategies. Regardless of those differences, the Visegrad countries recently exhibit a similar trend: persistently elevated food inflation. What is particularly striking is that food inflation in all four countries has exceeded both the average for the European Union and, according to the Food and Agriculture Organization (FAO, 2025) data, the global food price benchmark.

Figure 1 shows this trend using monthly data from January 2009 to May 2025, capturing the post-Euro adoption period for Slovakia and allowing for a comparative analysis of monetary policy divergence within the V4 region. As illustrated, while food inflation remained relatively stable and close to the European Union average (FOOD_



Note: FOOD denotes the Food Price Index published by Eurostat. WFPI refers to the World Food Price Index published by the FAO. Food inflation index for HICP indices are expressed with base year 2015 = 100; WFPI is expressed with base period 2014–2016 = 100

Source: Prepared by author

Figure 1: Evolution of the food inflation in the Visegrad Group and other regions.

EU27) until 2020, the post-pandemic period shows a sharp divergence. The most significant increase was recorded in Hungary, reaching levels significantly above both the EU and global averages (WFPI). By mid-2023, WFPI began to decline, yet V4 countries continued to face elevated food inflation. These differences raise a crucial question: can monetary policy and other related macroeconomic indicators be held responsible for the observed differences in food price inflation across these countries?

Understanding food inflation has become increasingly important in the post-pandemic period, when price pressures intensified globally and exposed vulnerabilities in food supply chains. The Visegrad region is a particularly relevant case because it combines structural similarities (geography, economic openness, integration into EU markets) with substantial differences in monetary policy frameworks. These differences create a unique natural setting for examining how food prices respond to monetary tightening under distinct institutional regimes. Moreover, food inflation carries strong social and political implications in the V4 countries, where household expenditure shares on food remain above the EU average, making price volatility more damaging to welfare and consumption stability. Despite the relevance of the topic, systematic evidence for the Visegrad countries is surprisingly limited. Existing research has largely focused on developing regions or larger emerging economies, while central Europe has received significantly less attention (Fertő and Bareith, 2024), even though the region has experienced some of the sharpest food price increases in the EU. Therefore, expanding empirical research in this area is essential for understanding both country-specific inflation patterns and broader regional trends.

While conventional OLS regression provides insights into the average relationship between monetary policy and food inflation, such an approach may overlook important distributional patterns. According to several authors (e.g., Iddrisu and Alagidede, 2021; Ali et al., 2022), quantile regression offers a clear advantage: it allows the analysis to capture heterogeneous effects along the entire distribution of food inflation, distinguishing periods of low, moderate, and high inflation. This study addresses the following research question: How does monetary policy affect food inflation in the Visegrad countries across different inflation regimes? This method has been increasingly used in mentioned empirical studies and provides additional insights beyond OLS by revealing whether monetary policy has stronger effects in the upper tail of inflation, where pressures tend to be most persistent. Given the substantial variability and cross-country differences in food inflation within the Visegrad Group, quantile regression is particularly well suited to uncover the asymmetric nature of monetary transmission in this region.

Literature review

Understanding the relationship between monetary policy and food inflation has its relevance stems not only from growing inflationary pressures in the examined region (e.g., Fertő and Bareith, 2024), but also other studies underscore its significance worldwide. For example, Kargbo (2000, 2005), Asfaha and Jooste (2007) or Iddrisu and Alagidede (2020) have examined this relationship in several African economies, or have been focused on Asian countries (Ali et al., 2025), particularly India (e.g., Anand et al., 2014; Samal et al., 2022), Iran (Hatami et al., 2022), Pakistan (e.g., Ali et al. 2022) and in the United States (Frankel, 2006 or Alam and Gilbert, 2017). A more recent contribution by Sami and Makun (2024) examines how monetary policy can help stabilize food inflation across five major emerging economies (Brazil, China, India, Russia, and South Africa, also known as BRICS). These studies converge on the central theme of understanding how food inflation reacts to monetary and macroeconomic conditions, especially in developing or transforming economies. Despite this, there is a noticeable lack of empirical research examining the nexus between monetary policy and food inflation in European contexts. Notably, Bhattacharya and Jain (2020) have already confirmed that food inflation serves as a leading indicator for the policy rate in Hungary. Given the heterogeneity in monetary frameworks and economic structures across the Visegrad Group countries, this finding motivates a closer examination of the relationship between monetary policy and food inflation in the region.

Many of the mentioned studies (e.g., Iddrisu and Alagidede, 2021; Ali et al., 2022 or Sami and Makun, 2024) employ quantile regression to uncover asymmetric effects across different levels of food inflation, emphasizing the importance of context-specific and distribution-sensitive policy responses. Applying quantile regression to the Visegrad countries allows us to uncover whether the effects of monetary policy on food inflation vary across the distribution of food price changes, and whether monetary policy has a stronger effect during periods of low or high food inflation. While the core aim of the research is to examine the relationship between monetary policy and food inflation, the analysis also considers other key macroeconomic variables, such as exchange rates, global food prices, industry, transport and energy prices that can be identified in the literature as relevant determinants of food price movements.

Building on the growing body of literature, this study employs the ordinary least squares (OLS) and quantile regression (QR) to examine the relationship between monetary policy and food inflation in the Visegrad Group countries. Ali et al. (2022) argue that conventional econometric methods such as VAR and ARDL fail to capture the tail behaviour of food inflation, which is critical in countries like Pakistan, where food inflation disproportionately affects vulnerable households. Their findings indicate that contractionary monetary policy and rising transportation costs significantly raise food inflation across all quantiles. Similarly, Iddrisu and Alagidede (2021) provide empirical evidence from Ghana showing that tight monetary policy destabilizes food prices especially in the lower to middle parts of the food inflation distribution, suggesting asymmetric effects not visible in average-based models. Sami and Makun (2024) extend this perspective to a panel of BRICS countries and confirm a causal link between monetary policy and food inflation. Their analysis highlights the heterogeneous effects of external shocks, such as energy prices and exchange rates, on domestic food prices, while also noting that contractionary monetary policy tends to reduce food inflation when macroeconomic coordination is effective. These findings support the use of econometric methods, such as OLS and especially quantile regression, which enable a more nuanced understanding of how monetary policy interacts with food inflation across different inflationary environments.

Materials and methods

The empirical strategy of this study is designed to systematically examine how monetary policy influences food inflation in the Visegrad countries. The research design follows a structured sequence: (i) identifying the relevant macroeconomic channels through which food prices adjust, (ii) constructing country-specific datasets that reflect the institutional differences in monetary policy frameworks, and (iii) estimating econometric models capable of capturing both average and distributional effects.

The analysis is conducted separately for each Visegrad country, rather than as a pooled panel, because the monetary policy frameworks differ fundamentally across the region. Czechia, Hungary, and Poland operate independent inflation-targeting regimes, where domestic central banks adjust policy rates in response to local economic conditions. In contrast, Slovakia has been part of the euro area since 2009, meaning its monetary stance is determined at the supranational level by the European Central Bank. The data employed in this study consists of monthly observations from January 2009 to May 2025 and the choice of variables is grounded in previous studies. The variables used in this analysis include food inflation (FOOD), monetary policy rate (MPR),world food price index (WFPI), exchange rate (ER), production development (INDUSTRY), harmonised index of consumer prices for transport (TRANSPORT) and energy (ENERGY) as a variable for the robustness checks. The selection of these variables is as follows:

- FOOD serves as the dependent variable in all estimated models. It is measured using the Harmonised Index of Consumer Prices (HICP) for food, sourced from Eurostat. As the HICP follows a unified and internationally comparable methodology across EU member states, it enables consistent cross-country comparison of food price developments within the Visegrad Group. The use of this indicator as the main measure of food inflation is in line with recent empirical studies on food price behaviour (e.g., Ali et al., 2022).
- MPR represents the key interest rate set by each country's central bank and is one of the key explanatory variables in the analysis. The data are sourced from the Bank for International Settlements (BIS). Specifically, for the Czech Republic, it corresponds to the two-week repo rate set by the Czech National Bank. In Slovakia is the rate on the deposit facility of the Eurosystem. For Hungary and Poland, the MPR is represented by the base rate determined by the Magyar Nemzeti Bank and the National Bank of Poland, respectively. To reduce simultaneity bias, the monetary policy rate enters the model with a three-month lag (MPR_{t-3}), reflecting the short-run transmission window documented by Fertő and Bareith (2024).
- WFPI refers to the World Food Price Index by the Food and Agriculture Organization of the United Nations on a monthly basis. Although Ali et al. (2022) analyse this relationship in Pakistan and Samal et al. (2022) in India, both studies highlight a general mechanism that is relevant for any open economy: domestic food prices tend to co-move with global commodity prices due to import dependence, integrated supply chains and exposure to global shocks.
- ER is expressed as the number of units of local currency per one euro based on monthly data. It captures the exchange rate between the euro and the local currency of each Visegrad country. For Slovakia, this variable was excluded from the analysis, as Slovakia has used the euro as its official currency throughout the examined period. As emphasised by Fertő and Bareith (2024), in small open economies, the strength of the national currency plays a crucial role in shaping inflation outcomes, given their high exposure to international markets and imported price pressures.
- INDUSTRY captures domestic production dynamics and is measured using the industrial production index, reflecting changes in output over the previous three months, sourced from Eurostat. Ali et al. (2022) employ a similar indicator, changes in large-scale manufacturing output, to show that fluctuations in domestic production capacity can exert meaningful effects on food inflation, particularly in emerging economies.
- TRANSPORT is derived from the HICP sub-index for transport from Eurostat. Ali et al. (2022) demonstrate that rising transport costs exert a significant upward pressure on food inflation across multiple quantiles.
- ENERGY also based on Eurostat HICP sub-indices, is used as a control variable in robustness checks. Energy prices can significantly affect both production and transportation costs, potentially making this variable even more influential than transport inflation in certain contexts.

The expected effects of the selected explanatory variables on food inflation are guided by previous empirical work, particularly the methodologies and findings of the mentioned studies. Based on their findings, we anticipate that an increase in the MPR will be associated with rising food inflation, as consistently observed in both studies. The impact of the WFPI and ER is assumed to be variable and possibly inconsistent across quantiles, given the mixed results reported in both studies. In terms of INDUSTRY, we expect a negative relationship with food inflation, while a positive relationship is expected for TRANSPORT, as highlighted by Ali et al. (2022). The data is collected from Eurostat

(2025), while the MPR is collected from BIS Data Portal (2025) and the WFPI data is collected from the FAO (2025).

A key methodological concern in estimating the relationship between monetary policy and food inflation is the potential endogeneity of the policy rate. Because central banks adjust interest rates in response to inflation developments, contemporaneous regressions may suffer from reverse causality. To ensure that non-stationarity does not bias the results, all variables are tested for unit roots using the Phillips–Perron test. Variables that are non-stationary in levels are transformed into first differences.

As a baseline approach, the study estimates a standard linear regression model (OLS) by using 1st differences (Δ):

$$\Delta FOOD_t = \beta_0 + \beta_1 MPR_{t-3} + \beta_2 \Delta WFPI + \beta_3 \Delta ER + \beta_4 \Delta INDUSTRY + \beta_5 \Delta TRANSPORT + \varepsilon_t \quad (1)$$

The dependent variable $\Delta FOOD_t$ therefore reflects the monthly change in food inflation, while the explanatory variables capture monthly changes in global food prices, exchange rates, industrial production, and transport inflation. The monetary policy rate is included in its lagged form (MPR_{t-3}) to account for the delayed transmission of monetary policy. Because all macroeconomic variables enter the model in first differences and the monetary policy rate is included with a three-month lag, the estimated coefficients should be interpreted as short-run effects. They capture how changes in the policy rate three months earlier and contemporaneous changes in the control variables are associated with the current month-to-month change in food inflation, rather than long-run equilibrium relationships. In the robustness section, the ΔTRANSPORT variable in equation (1) is substituted with ΔENERGY, following evidence that energy price inflation can significantly affect food price developments. This adjustment has been economically justified according to Moessner (2025) who states that energy inflation has a significantly positive effect on food inflation, considering a panel of 36 OECD countries. Energy costs influence both agricultural production and food distribution, and may therefore capture broader input-cost pressures than transport inflation alone. This alternative specification allows us to evaluate whether the impact of monetary policy remains stable when the model incorporates a more comprehensive cost-side indicator.

Then, we apply quantile regression to examine how monetary policy and selected control variables affect various quantiles of food inflation in the Visegrad group countries.This method allows us to examine the relationships between the dependent and independent variables across different points of the conditional distribution, specifically the 20th, 40th, 60th, and 80th quantiles. Such an approach enables the analysis of distributional heterogeneity that cannot be captured through mean-based regression techniques. We consider quantile regression framework originally developed by Koenker and Bassett (1978), leading by the following equations:

$$f_t = x'_t \beta + \mu_t \quad (2)$$

$$E\,(f_t \mid x_t) = x'_t \beta \quad (3)$$

$$Q_{ft}(\tau \mid x_t) = x'_t \beta_\tau \quad (4)$$

$$\beta_\tau = \beta + \vartheta F^{-1}(\tau) \quad (5)$$

In this context, it is assumed that the error terms μ_t are independently and identically distributed. The variable F denotes the cumulative distribution function of the error term μ_t, while ϑ represents a constant. The number of quantiles is represented by τ, while $Q_{ft}(\tau|x_t)$) refers to the conditional quantile function of food inflation given the covariates x_t. We use a dataset with 197 observations, dividing the dependent variable into quantiles (20th, 40th, 60th, 80th) to ensure that each quantile contains a sufficient number of observations (Ali et al., 2022), thus avoiding issues related to degrees of freedom and estimation accuracy. β_τ reflects the effect of the explanatory variables at a specific quantile of the distribution, capturing the marginal impact of the regressors.

The estimation of the quantile coefficients (parameters in equations 1 to 4) involves minimizing the following loss function:

$$\max_{\beta_\tau \in \Re p} \sum_{t=1}^{T} \rho_\tau \,(f_t - x'_t \beta), \quad (6)$$

$$\rho_\tau(\mu) = \mu(\tau - I(\mu < 0)) \quad (7)$$

where an indicator functionI that equals 1 when μ < 0_μ< 0, and 0 otherwise. In contrast to mean regression techniques, quantile regression minimizes the sum of the absolute residuals, making it more robust to outliers and non-normal error distributions (Iddrisu and Alagidede, 2020 and Ali et al., 2022).

Results and discussion

Firstly, to better understand the variable behaviour across the Visegrad countries, we begin by providing a summary of the descriptive statistics for the key indicators included in our analysis. Table 1 presents the descriptive statistics of the variables used in the analysis for the Visegrad Group countries: Czechia, Hungary, Poland, and Slovakia.

In Table 1, FOOD shows notable variation across the countries. The highest mean value is observed in Hungary (120.82), followed by Poland (115.35), Slovakia (112.41), and Czechia (109.28). The volatility, as captured by the standard deviation, is also greatest in Hungary (38.93), reflecting stronger fluctuations in food prices over time. The maximum food inflation index reaches 212.78 in Hungary, significantly higher than in the other countries. MPR also varies considerably among the countries, reflecting differences in policy frameworks and responses to inflation. Hungary again shows the highest average rate (4.37%)

Statistics		Czechia	Hungary	Poland	Slovakia
FOOD	Min:	79.3	80.17	87.6	84.22
	Median:	103.8	102.37	104.6	101.59
	Mean:	109.28	120.82	115.35	112.41
	Max:	154.8	212.78	176.4	170.48
	STD:	21.27	38.93	24.59	24.65
MPR	Min:	0.05	0.6	0.1	0
	Median:	0.75	3.2	2.5	0.25
	Mean:	1.72	4.37	3.02	0.9
	Max:	7	13	6.75	4.5
	STD:	2.24	3.72	1.94	1.31
WFPI	Min:	84.3			
	Median:	114.1			
	Mean:	111.5			
	Max:	160.2			
	STD:	17.13			
ER	Min:	23.44	265.29	3.88	x
	Median:	25.65	311.96	4.29	x
	Mean:	25.78	325.97	4.3	x
	Max:	28.46	418.31	4.8	x
	STD:	1.06	41.19	0.2	x
INDUSTRY	Min:	-62.9	-59.1	-53.3	-76.3
	Median:	5.1	0.8	-0.4	7.6
	Mean:	2.04	-5.18	-2.3	6.41
	Max:	26.1	24.7	16.1	56.3
	STD:	13.73	18.47	10.25	20.71
TRANSPORT	Min:	92.6	78.24	87	94.7
	Median:	104.1	105.82	104.6	106.4
	Mean:	109.63	111.67	109.52	109.09
	Max:	141.6	158.75	141.1	138.39
	STD:	13.93	20.38	13.69	12.49
ENERGY	Min:	86.8	87.09	81.7	90.68
	Median:	103.7	108.36	103.7	101.12
	Mean:	114.72	115.97	113.55	105.07
	Max:	185.5	173.43	173.1	130.43
	STD:	28.61	21.62	25.63	11.62

Source: Prepared by author

Table 1: Descriptive statistics.

and maximum value (13.00%), suggesting a more aggressive monetary stance compared to its regional peers. In contrast, Slovakia, which is part of the euro area, records the lowest average policy rate (0.90%), aligned with the historically low policy rates set by the European Central Bank. Czechia and Poland fall in between, with average rates of 1.72% and 3.02%, respectively. Regarding INDUSTRY, Slovakia exhibits the highest mean growth (6.41) and a maximum value of 56.30, while Hungary and Poland show negative average values, indicating structural weaknesses or downturns during parts of the sample period. ENERGY and TRANSPORT indices show relatively higher variability in Hungary, reflecting volatility in these sectors. This descriptive statistics reveal strong heterogeneity in food inflation, reflected in wide min–max ranges, high volatility, and cross-country differences. This indicates that food inflation behaves differently across various inflation regimes, which motivates the use of quantile regression to capture these heterogeneous effects.

Then, to assess the stationarity of the data, we applied the Phillips and Perron (1988) test similarly as Ali et al. (2022), Samal et al. (2022) or Fertő and Bareith (2024). The results in Table 2 indicate that most variables are non-stationary in levels but become stationary after first differencing.

To address concerns regarding endogeneity and reverse causality, we perform Granger causality tests separately for each Visegrad country and show those results in Table 3. The tests are estimated on country-specific series, as monetary policy rates differ fundamentally across countries and cannot be meaningfully pooled into a panel structure. For each country, we evaluate whether the policy rate Granger-causes food inflation and whether food inflation Granger-causes the policy rate. This approach allows us to assess the direction of causality within each monetary regime.

The Granger causality tests reveal a clear and consistent pattern across the three countries with independent monetary policy (Czechia, Hungary, and Poland). In all three cases, changes in the monetary policy rate Granger-cause changes in food inflation. This indicates that monetary policy contains predictive information for subsequent food price dynamics. In contrast, food inflation does not Granger-cause the policy rate in any of these countries, suggesting the absence of reverse causality and mitigating concerns about endogeneity. For Slovakia, although the underlying data on food inflation and the ECB policy rate are complete, the policy rate exhibits very limited month-to-month variation over most of the sample. As a result, the first-differenced policy rate contains insufficient variation for estimating a VAR model

Variable	Czechia		Hungary		Poland		Slovakia	
	Level	1st diff	Level	1st diff	Level	1st diff	Level	1st diff
FOOD	-4.73	-166.74***	-1.79	-74.60***	-0.3	-101.82***	-1.35	-128.90***
MPR	-5.92	-149.8***	-4.39	-54.94***	-3.24	-54.941***	-4.37	-172.31***
WFPI	-6.26	-119***	-6.26	-119***	-6.26	-119***	-6.26	-119***
ER	-8.99	-172.88***	-17.51	-167.68***	-19.24*	-157.65***	x	x
INDUSTRY	-30.57***	x	-13.1	-217.45***	-24.95**	-137.48***	-71.10***	x
TRANSPORT	-3.75	-117.13***	-3.03	-110.72***	-4.13	-176.68***	-4.59	-144.89***
ENERGY	-3.92	-204.46***	-3.91	-129.77***	-1.33	-212.34***	-2.71	-191.76***

Source: Prepared by author

Table 2: Stationary test.

Country	Direction of Causality	F-Statistic	p-value	Interpretation
Czechia (CZ)	ΔMPR → ΔFOOD	3.3756	0.0030	Significant causality
	ΔFOOD →Δ MPR	1.2692	0.2708	No causality
Hungary (HU)	ΔMPR → ΔFOOD	5.9981	<0.001	Significant causality
	ΔFOOD →Δ MPR	1.5216	0.1700	No causality
Poland (PL)	ΔMPR → ΔFOOD	4.6948	<0.001	Significant causality
	ΔFOOD → ΔMPR	1.3225	0.2461	No causality
Slovakia (SK)	ΔMPR →ΔFOOD	—	—	Not estimable
	ΔFOOD → ΔMPR	—	—	Not estimable

Source: Prepared by author

Table 3: Granger causality test results for country-specific models.

or conducting Granger causality tests. This is consistent with Slovakia's membership in the euro area, where monetary policy is not determined at the national level. While causality cannot be assessed, quantile regression could evaluate how the common monetary stance of the euro area and other macroeconomic factors are associated with different parts of the distribution of Slovak food inflation.

While Granger causality tests do not identify structural causal effects, they help to mitigate endogeneity concerns by clarifying the temporal ordering between monetary policy and food prices. The finding that changes in the policy rate Granger-cause food inflation, but not vice versa, supports the use of lagged policy rates in the subsequent regressions and reduces the risk that our estimated coefficients merely capture the reaction of central banks to inflation. The choice of a three-month lag for the policy rate (MPR_{t-3}) is motivated by the literature on food price dynamics (e.g., Fertő and Bareith, 2024), which documents that short-run adjustments occur mainly within one to three months after a monetary shock.

Table 4 presents the OLS and quantile regression estimates for food inflation in Czechia across the 20th, 40th, 60th, and 80th quantiles. The results for Czechia reveal that the determinants of food inflation operate asymmetrically across the distribution, underscoring the relevance of quantile regression for capturing heterogeneous price dynamics. The monetary policy rate shows a positive and statistically significant effect only in the upper tail (80th quantile), suggesting that interest rate changes matter primarily during high-inflation period. Once ENERGY is included (Table 8 in Appendix A), the policy rate becomes insignificant across all quantiles, suggesting that short-run monetary transmission is weak in an environment where energy and production costs dominate price formation. This aligns with recent evidence for small open European economies, where energy price shocks have been shown to transmit rapidly into food prices, often overshadowing the direct impact of monetary policy.

The results for Hungary (Table 5) reveal a clear and economically meaningful pattern. The monetary policy rate shows a stronger and more systematic relationship with food inflation than in any other Visegrad country. In the baseline specification, MPR_{t-3} is statistically significant in the 20th and 80th quantiles, with a negative effect at the lower tail and a positive effect in periods of elevated food inflation. Tighter monetary policy dampens food prices when inflation is low but amplifies them in high-inflation regimes. The negative effect in the lower tail suggests that when food inflation is low, tighter monetary policy helps reduce price pressures through weaker demand. In contrast, the positive and significant effect at the 80th quantile indicates that during high-inflation periods, interest rate increases can push food prices up. This finding is consistent with the evidence reported by Ali et al. (2022). INDUSTRY is negative and consistently significant in the OLS, 60th, and 80th quantiles, suggesting that stronger industrial output helps ease food price pressures, particularly

Variable	OLS		20th Quantile		40th Quantile		60th Quantile		80th Quantile	
Intercept	0.22323	**	-0.62957	***	0.04998		0.41500	***	1.04032	***
	(0.12672)		(0.13088)		(0.13762)		(0.09827)		(0.21782)	
MPR_{t-3}	0.07034		-0.04918		-0.04493		0.07242		0.23976	**
	(0.04478)		(0.05908)		(0.00412)		(0.06568)		(0.12048)	
ΔWFPI	-0.04104		0.03831		0.04952		-0.02785		-0.13674	*
	(0.03759)		(0.05444)		(0.53763)		(0.03940)		(0.07216)	
ΔER	-0.00467		0.69187		0.44112		0.41163		0.25720	
	(0.35007)		(0.54236)		(0.42513)		(0.26384)		(0.50820)	
ΔINDUSTRY	-0.02166		-0.02613		-0.02064		-0.01096		-0.01652	
	(0.01435)		(0.01727)		(0.01591)		(0.01528)		(0.02074)	
ΔTRANSPORT	0.15399	**	-0.03515		-0.12642		0.05442		0.38146	**
	(0.08602)		(0.12328)		(0.10297)		(0.10597)		(0.15998)	
R-squared	0.0417									
F-statistic	1.636									
Observations	194		194		194		194		194	

Note: *, **, *** represents significant at 10%, 5% and 1%; value in parenthesis () represents the standard error
Source: Prepared by author

Table 4: Quantile regression estimates for Czechia.

Variable	OLS		20th Quantile		40th Quantile		60th Quantile		80th Quantile	
Intercept	0.50813	**	-0.11267		0.21703	*	0.48900	***	0.95108	***
	(0.20682)		(0.16434)		(0.12643)		(0.18695)		(0.23572)	
MPR_{t-3}	0.02246		-0.11990	**	-0.00754		0.06489		0.12039	***
	(0.03607)		(0.05585)		(0.03583)		(0.04790)		(0.05567)	
ΔWFPI	-0.03750		0.00919		0.03971		0.09556		0.02826	
	(0.04582)		(0.06889)		(0.05816)		(0.05902)		(0.09013)	
ΔER	-0.01626		-0.00252		-0.03530		-0.04254		-0.00863	
	(0.02540)		(0.04278)		(0.02836)		(0.02985)		(0.03434)	
ΔINDUSTRY	-0.04547	**	-0.01488		-0.03076		-0.04303	**	-0.05807	**
	(0.02101)		(0.02194)		(0.02137)		(0.02069)		(0.02593)	
ΔTRANSPORT	0.13054	**	0.02491		0.09389		0.09893		0.12442	
	(0.06231)		(0.08786)		(0.06878)		(0.07614)		(0.09171)	
R-squared	0.0493									
F-statistic	1.95	*								
Observations	194		194		194		194		194	

Note: *, **, *** represents significant at 10%, 5% and 1%; value in parenthesis () represents the standard error
Source: Prepared by author

Table 5: Quantile regression estimates for Hungary.

in the middle and upper parts of the distribution. Other variables, such as WFPI, ER, and TRANSPORT, are generally insignificant, which implies that short-run domestic food price changes in Hungary are driven more by local cost and supply conditions than by immediate global or exchange-rate shocks. The robustness check (Table 9 in Appendix A), strengthens this interpretation. Taken together, the results highlight that Hungary's food inflation is heavily shaped by domestic production and cost factors, while monetary tightening tends to raise food prices during periods of high inflation rather than dampen them.

The results for Poland (Table 6) indicate a strong and fairly consistent role of monetary policy across the higher parts of the food inflation distribution. This pattern suggests that tighter monetary policy is associated with higher monthly food inflation particularly during moderate-to-high inflation regimes. Industrial production (ΔINDUSTRY) shows a negative and significant effect in the OLS, 20th, and 40th quantiles, implying that stronger domestic production capacity helps dampen food price pressures, but primarily in lower-inflation environments. Other variables, including global food prices (ΔWFPI) and the exchange rate (ΔER) remain statistically insignificant across quantiles. This suggests that external price pressures may transmit to Polish food inflation more gradually, beyond the monthly horizon examined here. When ENERGY is introduced in the robustness specification (Table 10 in Appendix A), the results for Poland become even more consistent. The monetary policy coefficient remains positive and significant in the higher quantiles, while ENERGY emerges as a strong and statistically significant driver of food inflation across almost the entire distribution.

Slovakia (Table 7) shows no measurable impact of monetary policy on food inflation across any part of the distribution, reflecting the fact that the country does not set its own policy rate within the euro area. Most macroeconomic controls also remain insignificant in the baseline model, indicating limited short-term pass-through of external and structural shocks. By contrast, ENERGY (Table 11 in Appendix A) becomes strongly significant across nearly all quantiles in the robustness test, suggesting that cost-push factors—rather than monetary policy—drive monthly food price fluctuations in Slovakia.

The quantile regression results across Czechia, Hungary, Poland, and Slovakia show that the transmission of monetary policy to food inflation is heterogeneous and concentrated mainly in higher-inflation regimes. In Poland, and to a lesser extent in Hungary, the three-month lagged policy rate is positively and significantly associated with food inflation in the middle and upper quantiles, indicating that monetary tightening amplifies food price pressures when inflation is already high. In Czechia, the effect appears only in the upper tail of the distribution, and in Slovakia the policy rate is insignificant at all quantiles, consistent with the absence of a national monetary policy under the euro area framework.

Variable	OLS		20th Quantile		40th Quantile		60th Quantile		80th Quantile	
Intercept	0.10274		-0.47450	**	-0.17857		0.12748		0.80110	***
	(0.16586)		(0.19826)		(0.16959)		(0.16502)		(0.23843)	
MPR_{t-3}	0.11028	**	0.01881		0.11078	**	0.17030	**	0.14490	*
	(0.04670)		(0.08531)		(0.05265)		(0.05444)		(0.08663)	
ΔWFPI	-0.00370		-0.01562		0.03446		0.01850		0.00695	
	(0.03261)		(0.05859)		(0.04252)		(0.03113)		(0.05426)	
ΔER	-0.48457		1.47467		-0.32335		-0.10649		0.39600	
	(1.38370)		(2.01541)		(1.61178)		(1.39118)		(2.08754)	
ΔINDUSTRY	-0.04344	**	-0.04375	*	-0.03944	**	-0.02774		0.00307	
	(0.02047)		(0.02634)		(0.01802)		(0.01935)		(0.03268)	
ΔTRANSPORT	0.09685	*	0.07658		0.07488		0.13766	*	0.10392	
	(0.05369)		(0.11760)		(0.08046)		(0.07060)		(0.07924)	
R-squared	0.0609									
F-statistic	2.439	**								
Observations	194		194		194		194		194	

Note: *, **, *** represents significant at 10%, 5% and 1%; value in parenthesis () represents the standard error
Source: Prepared by author

Table 6: Quantile regression estimates for Poland.

Variable	**OLS**		**20th Quantile**		**40th Quantile**		**60th Quantile**		**80th Quantile**	
Intercept	**0.42151**	***	**-0.50605**	***	**0.04375**		**0.42084**	***	**1.15785**	***
	(0.10621)		**(0.10652)**		**(0.08724)**		**(0.11392)**		**(0.18079)**	
MPR_{t-3}	**-0.02426**		**-0.05371**		**-0.00571**		**0.00536**		**0.03874**	
	(0.06735)		**(0.09782)**		**(0.05869)**		**(0.08423)**		**(0.15322)**	
ΔWFPI	**-0.01117**		**0.01494**		**-0.00830**		**-0.03728**		**-0.04978**	
	(0.03099)		**(0.04947)**		**(0.03647)**		**(0.03792)**		**(0.05692)**	
ΔINDUSTRY	**-0.00281**		**-0.00734**		**-0.00520**		**-0.00125**		**-0.00864**	
	(0.00485)		**(0.00532)**		**(0.00499)**		**(0.00738)**		**(0.00730)**	
ΔTRANSPORT	**0.07749**		**-0.03408**		**0.06025**		**0.07469**		**0.26621**	*
	(0.05438)		**(0.09506)**		**(0.04760)**		**(0.08958)**		**(0.15133)**	
R-squared	**0.01277**									
F-statistic	**0.6112**									
Observations	**194**		**194**		**194**		**194**		**194**	

Note: *, **, *** represents significant at 10%, 5% and 1%; value in parenthesis () represents the standard error
Source: Prepared by author

Table 7: Quantile regression estimates for Slovakia.

Across countries, the significance of WFPI and ER is generally weak or inconsistent, implying that global food price shocks and exchange rate movements exert only limited short-run effects on monthly food inflation—a finding consistent with earlier evidence that their pass-through operates with longer lags (e.g., Iddrisu and Alagidede, 2020; Ali et al., 2022). The INDUSTRY variable shows a predominantly negative association with food inflation, particularly in Poland and Hungary, although statistical significance is confined to the lower and middle quantiles, where domestic production dynamics appear more stabilising. TRANSPORT inflation provides only weak and irregular explanatory power, with coefficients often insignificant across quantiles. In contrast, the robustness check (Appendix A) replacing TRANSPORT with ENERGY reveals a clearer and more stable positive effect of energy price inflation, especially in Poland and Slovakia. This pattern underscores the importance of cost-push channels, most notably energy-intensive inputs, in shaping food price dynamics in the region and confirms that supply-side factors can overshadow the direct effects of monetary policy in certain inflation environments.

Regarding to the Visegrad region, the results suggest that domestic monetary conditions matter primarily in countries with independent monetary policy and mainly during high-inflation episodes. The repeated insignificance of global food prices and exchange rates implies that their short-run pass-through is limited in monthly data, while cost-side pressures, particularly energy, play a more prominent role. The findings also show that in Slovakia, unlike the other V4 countries, energy inflation systematically shapes food price changes, whereas the interest rate channel is largely absent.

From a policy perspective, the results suggest that the short-run impact of monetary policy on food inflation in the Visegrad countries operates over a horizon of three months, with the strongest effects observed in higher-inflation regimes and in countries with independent monetary policy, most notably Hungary and Poland. This implies that central banks in these economies should take into account the delayed and asymmetric response of food prices when calibrating interest-rate decisions, particularly during periods of elevated inflation when monetary tightening appears to exert upward cost-push pressure rather than dampening food price dynamics. For Czechia, the monetary policy effect is confined to the upper quantile and disappears once energy inflation is included, suggesting that domestic food prices respond more strongly to cost-side shocks than to the policy rate itself. In Slovakia, where monetary policy is determined at the euro area level, national food inflation is largely unresponsive to the policy rate. Instead, the results indicate that domestic food prices are primarily driven by fundamental cost factors such as energy inflation, which remains consistently significant across quantiles. This highlights that, in the absence of national monetary tools, external price pressures and energy-related input costs dominate short-run food price developments.

Conclusion

The interaction between food price inflation and monetary policy is a complex and increasingly relevant issue and recently it has been highlighted by several authors such as Sami and Makun (2024), Fertő and Bareith (2024) or Ali et al. (2025).The main objective of this study was to examine how monetary policy affects food inflation, while simultaneously assessing the role of key fundamental determinants such as global food prices, exchange rates, industrial production, transport costs, and energy inflation, in the Visegrad Four countries – Czechia, Hungary, Poland, and Slovakia. To capture the heterogeneous effects of monetary and other economic variables across different levels of food price inflation, both OLS and quantile regression methods were applied. This dual approach allowed for a better understanding of the determinants of food inflation beyond average effects and revealed how relationships differ across the inflation distribution.

The OLS results provided an initial overview, indicating that monetary policy rates are positively associated with food inflation in several V4 countries. However, the quantile regressions revealed a more nuanced picture: the effect of monetary policy is asymmetric and becomes statistically significant primarily in the upper quantiles of the inflation distribution. The model is estimated in first differences, and the policy rate enters with a three-month lag, therefore these results reflect only short-run, month-to-month dynamics. Within this short-run horizon, monetary transmission to food prices is most evident during periods of heightened inflation and in countries with independent monetary policy frameworks. This pattern is especially pronounced in Hungary and Poland, where MPR_{t-3} exhibits strong positive effects during periods of elevated food inflation. In Czechia, the effect is visible only at the upper tail, and in Slovakia, where monetary policy is determined at the euro area level, the policy rate is largely insignificant.

From an economic and policy perspective, the results highlight important limitations in the ability of monetary policy to stabilise food inflation. Although interest rate adjustments influence price dynamics at the margin, food inflation in the V4 countries is strongly shaped by supply-side and cost-push factors. This is underscored by the robustness checks, where the inclusion of energy inflation substantially improves model fit and reveals a consistently significant positive relationship with food prices across countries, most notably in Slovakia. Energy costs, global commodity price movements, and exchange rate fluctuations appear to exert larger and more persistent effects than short-term monetary policy measures. These insights point to the need for a multidimensional policy framework. While monetary policy can help anchor inflation expectations and dampen demand-driven pressures, it is insufficient to address structural and external drivers of food inflation.

In conclusion, the study provides empirical evidence that the impact of monetary policy on food inflation in the V4 countries is real but conditional: it is visible primarily in high-inflation regimes and interacts strongly with underlying structural and external factors. Quantile regression proves particularly valuable in uncovering these heterogeneous effects, offering a richer and more policy-relevant understanding of food price dynamics than traditional OLS estimates. The robustness results reinforce the central role of energy costs, indicating that future food inflation dynamics in the region will depend not only on monetary tightening cycles but also on global cost conditions and domestic structural adjustment.

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