



Dynamic vertical price transmission in the wheat–flour–bread market: Evidence from Bosnia and Herzegovina

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ABSTRACT

Price transmission in staple-food markets is particularly important in small, import-dependent economies, yet evidence for Bosnia and Herzegovina remains scarce. Using 44 quarterly observations from 2014 Q1 to 2024 Q4, we examine long-run co-movement, short-run adjustment, and possible asymmetric responses using a Vector Error-Correction framework supplemented by impulse-response analysis, variance decomposition, and auxiliary asymmetric error-correction models. Conventional Johansen inference supports a single cointegrating relationship, but this result is not robust to small-sample adjustment and is therefore interpreted cautiously. Conditional on the baseline rank-one specification, the strongest error-correction loading occurs in the flour consumer-price equation. Impulse-response point estimates are compatible with delayed and attenuated responses of flour and bread prices to wheat-price innovations. However, neither downstream response is statistically significant under the baseline ordering. Conditional on the maintained rank-one specification, HCl-robust Wald inference indicates asymmetric equilibrium correction in the bread consumer-price equation, while short-run asymmetry with respect to positive and negative wheat-price changes is only marginally supported at the 10% level. Overall, the findings suggest economically relevant dynamic co-movement among the observed price series, but provide limited support for strong claims of causal or stage-specific transmission. The results underline the need for improved market monitoring and more detailed import, cost, and stage-specific price data to support food-market policy in Bosnia and Herzegovina.

1. Introduction

Volatility in staple food prices poses significant challenges to economic stability, household welfare, and food security, particularly in small and developing countries [1–3]. Agricultural prices are influenced by interconnected factors, including market fundamentals, energy costs, exchange-rate dynamics, weather shocks, natural disasters, and policy interventions [4]. Monitoring price trends for essential agricultural commodities such as wheat has become increasingly important amid rising input costs and growing pressure on the global food system [5]. Recent disruptions associated with the COVID-19 pandemic and the conflict in Ukraine have intensified food inflation, deepening

vulnerabilities in developing economies and reversing progress on poverty and food security [6,7]. In the European Union, bread prices were, on average, 18% higher in August 2022 than one year earlier [8], while the FAO Cereal Price Index increased by approximately 33% between January 2020 and January 2022 [9]. Western Balkan countries experienced these effects acutely. Many countries in the region rely on wheat imports and other production inputs from Ukraine and Russia, exposing them to trade disruptions [10]. In March 2022, for example, rising global wheat prices prompted Serbia to impose a temporary ban on wheat and maize exports to stabilize its domestic market [9]. These developments underscore the need to understand how commodity price shocks are reflected in wheat, flour, and bread prices.

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Examining how price shocks are transmitted across the agri-food chain is important for understanding the implications of food price volatility. This process, known as *vertical price transmission*, describes the relationship between price movements at different points in the chain and provides insight into the speed, magnitude, and direction of price adjustment. In theory, the speed and completeness of pass-through may provide information about market functioning [11]. However, numerous studies reveal that price transmission is influenced by multiple factors, including transport costs, market power of certain actors within supply chains, economies of scale, product characteristics, exchange rates, and policy interventions [12]. Moreover, long-run price co-movement, the magnitude of pass-through, short-run adjustment, asymmetric responses, and market power are analytically distinct concepts. A weak or incomplete response of consumer prices to changes in agricultural commodity prices does not necessarily indicate market power, because raw materials represent only one component of the final product price. Labour, energy, transport, packaging, taxes, inventories, and other processing and distribution costs may also affect the magnitude and timing of price adjustments. Market power is therefore only one of several potential mechanisms that may lead to partial, delayed, or asymmetric price transmission [13]. Under certain market conditions, particularly in highly concentrated agri-food chains, such asymmetries may be associated with market inefficiencies and welfare distortions [14,15]. Nevertheless, evidence of delayed or asymmetric price adjustments should be interpreted with caution in the absence of direct information on costs, margins, market concentration or the pricing practices of market participants.

These issues are particularly relevant to the wheat-based food market in Bosnia and Herzegovina. Domestic wheat production meets only part of national consumption, while the country relies heavily on imports from neighbouring and EU markets. At the same time, wheat, flour, and bread are strategically important for domestic food security and household consumption. Despite this importance, the dynamic relationships between domestic wheat producer prices and consumer prices of flour and bread have not been systematically examined in Bosnia and Herzegovina. Existing research has primarily focused on Serbia and other European markets, leaving an important empirical gap concerning price adjustment within the country's decentralized institutional and import-dependent market context. Based on the available literature, this study provides the first systematic assessment of long-run co-movement, short-run adjustment, and potential asymmetry among these price series in Bosnia and Herzegovina.

1.1. Vertical price transmission of the wheat-flour-bread chain

Empirical studies of vertical price transmission distinguish several dimensions of price adjustment. Cointegration indicates whether non-stationary price series share a stable long-run relationship, but does not by itself demonstrate complete pass-through or causality [16,17]. Conditional on the selected normalization, the estimated cointegrating coefficients describe the magnitude and direction of the long-run relationship, while error-correction coefficients indicate how individual prices respond to deviations from that relationship. Short-run coefficients and impulse-response functions describe the timing and persistence of dynamic price responses, while generalized impulse-response analysis provides an ordering-invariant alternative [18]. Formal threshold-cointegration approaches allow equilibrium adjustment to differ across regimes or according to the direction of disequilibrium [19,20]. These approaches should be distinguished from more parsimonious auxiliary asymmetric ECM specifications that test parameter equality without estimating a formal threshold-cointegration mechanism. None of these price-based approaches can, on its own, identify market power, which requires additional evidence on production costs, quantities, margins, market concentration, or the pricing behaviour of market participants.

Empirical research from Serbia illustrates how conclusions from

price-transmission analysis depend on both the econometric model and the available data. Djurić et al. [21] found that the 2007–2008 wheat export ban neither changed the prevailing transmission regime nor reduced domestic wheat prices. Djurić and Götz [22] and Djurić [23] complemented price-transmission models with evidence on production costs, margins, and consumer expenditure and reported increased downstream margins and consumer welfare losses during intervention periods. Importantly, their conclusions regarding profits, margins, and welfare were supported by additional cost and expenditure information rather than by aggregate price series alone.

Evidence from other European markets confirms that price transmission patterns vary across products, periods, and model specifications. Ricci et al. [24] analyzed vertical price transmission of the Italian pasta and bread chains by separating each chain into upstream farm-wholesale and downstream wholesale-retail relationships. They applied cointegration methods that allowed for unknown structural breaks. They found clear evidence of asymmetric transmission in the pasta chain, particularly during periods of rising durum wheat prices. In contrast, the results for the bread chain were less conclusive and varied across regimes. The authors related these differences to market structure and the declining share of raw material costs in final consumer prices. Bareith et al. [25] examined weekly wheat producer prices and consumer prices of wheat flour in Hungary from 2017 to 2024 using linear and nonlinear cointegration tests and NARDL models. Their baseline NARDL specification indicated relatively small, long-run transmission effects. At the same time, the overall Wald tests did not reject long- or short-run symmetry, although crisis-specific estimates differed between the COVID-19 and the Ukraine war periods. Their causality results also suggested a stronger relationship from consumer flour prices to producer wheat prices than from producer wheat prices to consumer flour prices. Together, these studies show that evidence on the magnitude, direction, and asymmetry of transmission is sensitive to the product chain, sample period, and econometric specification. Brümmer et al. [26] used a Markov-switching VECM to examine wheat and flour prices in Ukraine and identified four transmission regimes whose timing coincided with major political and economic events. In particular, the association between a high-uncertainty regime and discretionary market interventions suggested that policy responses to harvest fluctuations may have amplified market instability; however, this temporal correspondence did not establish a causal policy effect. Rumánková [27] applied adjusted VECM specifications and impulse-response analysis to three Czech wheat-based product chains. The results were mixed across individual price relationships, but asymmetric reactions predominated when the chains were considered as a whole, with price increases generally transmitted faster or more completely than price decreases. Adjustment to price shocks was slow, with a new equilibrium reached after approximately 50 months. However, the study discussed several possible explanations for these patterns and did not treat asymmetric price transmission alone as proof of market power. Blažková and Syrovatka [28] compared vertical price transmission in Czech wheat-flour and wheat-roll chains using farm-gate, processor, and consumer prices. They found inelastic transmission and differences in the magnitude, direction, and timing of adjustment across products and price relationships. Although the authors interpreted some results as consistent with stronger retail bargaining power, the available price data could not separately identify wholesale and retail effects, and asymmetric adjustment could also reflect policy intervention, information asymmetries, or inventory management. Using a different approach, Cehura and Jaghdani [29] investigated the market imperfections in the milling and baking sectors of the French and United Kingdom wheat-to-bread chains using firm-level finance data from the Amadeus database for 2006–2018. They estimated mark-up and mark-down models within a stochastic frontier framework and derived Lerner indicators of market imperfections in both input and output markets. The findings indicated market imperfection in the milling and baking sectors, especially in the French baking sector, while firm size

was not generally associated with greater market power. Compared with aggregate price models, this firm-level approach provided a more direct basis for examining mark-ups, marks-downs and bargaining power. Finally, Syrovátka and Lechanová [30] used dynamic regression models to estimate price transmission and derived demand elasticities in the Czech food-grain chain. They found strongly inelastic transmission from wheat and rye commodity prices to consumer prices of flour and bakery products, with average transmission elasticities of 0.1602 and 0.1067, respectively.

Given the limited sample of 44 quarterly observations, the empirical strategy prioritizes a parsimonious specification. A substantially more parameterized threshold-VECM, NARDL, or regime-switching model would further reduce the effective degrees of freedom and would be difficult to estimate reliably in this sample. The study, therefore, uses a linear VECM as the conditional baseline model and auxiliary asymmetric ECM specifications as robustness checks. The latter are interpreted as tests of parameter asymmetry rather than as formally estimated threshold-cointegration models. To the best of our knowledge, this study provides the first systematic analysis of the dynamic relationships between wheat producer prices and consumer prices of flour and bread in Bosnia and Herzegovina.

The reviewed studies demonstrate that conclusions about vertical price transmission depend on both econometric method and available data. VECM and related price-based models can identify long-run co-movement, short-run dynamics, equilibrium correction, and possible asymmetries, whereas conclusions about costs, margins, profits, market power, or consumer welfare require additional firm or sector-level evidence. The literature also shows that the magnitude, speed, and direction of transmission vary across products, countries, policy regimes, and model specifications. Nevertheless, research on the wheat-flour-bread chain remains limited in the Western Balkan region outside Serbia, while Koveljanić et al. [10] identified broader analytical gaps in the region's cereal sector. This gap is particularly relevant for Bosnia and Herzegovina, where dependence on imported wheat and exposure to external price shocks coexist with a decentralized institutional framework.

1.2. The wheat sector in Bosnia and Herzegovina

The agri-food products market in Bosnia and Herzegovina, including the wheat sector, is characterized by significant decentralization and the absence of a unified state food security policy. Instead, responsibilities are divided between the entity level and the Brčko District. Adopting the Law on Federal Commodity Reserves [31] established a commodity reserve system in which publicly owned stocks may be stored and managed in cooperation with private operators. Conversely, in the Republic of Srpska, the Directorate for Commodity Reserves was dissolved in 2017 and replaced by a new system regulated by the Law on Intervention Procurement [32], under which private operators own and store goods while the government performs a supervisory and intervention role. In both entities, the primary function of commodity reserves is market intervention, particularly for agricultural products with administratively set prices, with wheat occupying a central position. When market prices fall below the guaranteed threshold, entities may purchase surplus products to stabilize the market. However, such interventions are not aligned with European Union market practices and may distort the functioning of the free market [33]. These institutional and market mechanisms are essential for analysing vertical price transmission in Bosnia and Herzegovina, as state interventions can potentially affect the timing and predictability of price signals along the supply chain.

Wheat, flour, and bread form the backbone of the agri-food sector, underpinning both food security and economic stability. Despite its status as a key staple food, domestic wheat production is limited in volume and quality [34]. Between 2014 and 2024, the average annual gross production value of wheat amounted to approximately USD 55.34

million [35], accounting for an average of 3.40% of the total gross value of agricultural production (authors' calculations based on [35]). Wheat production is also characterised by a fragmented farm structure, 50% of farms have less than 2 ha, and the average farm size is approximately 2.60 ha [36,37]. During the same period, an average of 64,375 ha was harvested annually, producing approximately 271,435 tonnes of wheat at an average yield of 4.24 t/ha (Table 1).

Wheat production increased from 170,550 tonnes in 2014 to 279,273 tonnes in 2024, representing an increase of 63.75% between the two endpoints, although substantial annual fluctuations occurred throughout the period. Based on the authors' calculations using FAO-STAT Food Balance data, the average self-sufficiency ratio for 2014–2023 was 32.35%, indicating that domestic production covered approximately one third of the available domestic wheat and wheat products supply. Over the same period, the average import dependency ratio was 83.24%, demonstrating the country's substantial reliance on imported wheat and wheat products. Average per capita wheat consumption amounted to 190.13 kg, equivalent to approximately 133.09 kg of flour per person. Prior research conducted by Ostojić et al. [38], analysing the period from 2014 to 2018, similarly identified a persistent structural wheat deficit in Bosnia and Herzegovina, with an estimated import dependency ratio of 77.79%. The country's substantial dependence on imported wheat increases the exposure of domestic wheat producer prices and downstream consumer prices to external supply and price shocks. However, this exposure does not imply complete transmission to flour and bread prices, which also incorporate labour, energy, transport, processing, and distribution costs.

Bosnia and Herzegovina also exhibits substantial dependence on wheat flour imports. Between 2021 and 2024, the country imported an average of 43,937 tonnes of wheat flour annually from CEFTA countries, and an average of 7592 tonnes from the European Union, based on authors' calculations using data from the Foreign Trade Chamber of Bosnia and Herzegovina [39]. Meanwhile, the domestic grain-milling industry is contracting, marked by a 2.8% annual decline in the number of businesses and a 4.4% decrease in market size from 2019 to 2024 [40]. Dependence on both wheat and flour imports exposes multiple stages of the domestic supply chain to external price shocks. It may affect the observed adjustment between producer and consumer prices. Bread nevertheless remains a major staple food, with per capita consumption estimated at approximately 48 kg [41]. By 2030, the total market volume is expected to reach 163.70 million kilograms, with a projected volume growth of 2.1% in 2026 [41]. Consequently, changes in wheat prices are not expected to be transmitted fully or immediately to bread prices, even under competitive market conditions. Against this background, Fig. 1 presents the quarterly price series used in the empirical analysis for the period 2014 Q1–2024 Q4.

Fig. 1 shows distinct trajectories in the three observed price series over the study period. Wheat producer prices remained relatively stable during the first part of the sample, followed by a pronounced increase during 2021–2022 and a subsequent decline. Flour consumer prices

Table 1
Area and wheat production in Bosnia and Herzegovina 2014–2024.

Year	Area harvested (ha)	Production (t)	Average yield (t/ha)
2014	59,550	170,550	2.86
2015	60,407	213,015	3.53
2016	71,394	306,605	4.29
2017	70,046	292,368	4.17
2018	73,591	295,810	4.02
2019	68,965	264,769	3.84
2020	69,914	321,084	4.60
2021	65,649	314,382	4.79
2022	51,541	281,414	5.46
2023	57,903	246,515	4.31
2024	59,168	279,273	4.72
Average	64,375	271,435	4.24

Source [35]:



Fig. 1. Quarterly wheat producer prices and flour and bread consumer prices in Bosnia and Herzegovina, 2014 Q1–2024 Q4 (EUR/kg). Source: Agency for Statistics of Bosnia and Herzegovina

increased markedly from 2018 onward, rose further during 2021–2023, and then declined sharply in 2024. Bread consumer prices exhibited a comparatively smoother upward trajectory, with the strongest increase occurring after 2021. These descriptive patterns indicate substantial differences in the timing and magnitude of price movements across the three series. Still, they do not establish the direction or statistical significance of price transmission. Formal time-series analysis is therefore required to examine long-run co-movement, short-run dynamics, and adjustment patterns. High import dependency increases the country's exposure to external commodity shocks. At the same time, fragmented domestic production, a contracting milling industry, and decentralized food-market governance may affect how price changes are transmitted across the domestic market.

Against this background, the study evaluates three testable hypotheses:

- H1.** A long-run equilibrium relationship links wheat producer prices and consumer prices of flour and bread;
- H2.** Wheat-price innovations are associated with delayed and attenuated responses in flour and bread consumer prices, with a weaker

response expected for bread.

H3. Adjustment differs between positive and negative deviations from equilibrium or changes in wheat producer prices.

Using quarterly data for 2014–2024, the study applies a baseline VECM to examine long-run relationships and short-run adjustment and supplements it with asymmetric error-correction specifications and Wald tests. These methods allow long-run co-movement, dynamic response patterns, and asymmetric adjustment to be examined separately, but they do not identify market power, marketing margins, or welfare effects.

2. Methodology

This section describes the data, preliminary time-series tests, baseline VECM specification, and robustness analyses.

2.1. Data and sources

This study uses 44 quarterly observations covering the period from 2014 Q1 to 2024 Q4. The empirical system combines the national

producer price of wheat with national consumer prices of white wheat flour and wheat bread obtained from the Agency for Statistics of Bosnia and Herzegovina. The flour series corresponds to the BHAS product category “Wheat flour, white”, while the bread series corresponds to “Wheat bread”. Both consumer-price series are reported per kilogram. Accordingly, the three variables should not be interpreted as direct firm-level prices at consecutive producer, milling, and retail stages of the supply chain. The wheat series represents an upstream producer-price measure, whereas the flour and bread series are consumer-price measures that incorporate processing, distribution, and retail components. Consumer-price observations, originally available at a monthly frequency, were aggregated to quarterly frequency using arithmetic quarterly averages to match the frequency of the wheat producer-price series. The resulting balanced dataset contains 44 observations per variable and no missing values. Before analysis, all price series were converted to EUR per kilogram using the fixed exchange rate of EUR 1 = BAM 1.95583. Wheat prices, originally reported per tonne, were divided by 1000 to obtain prices per kilogram before currency conversion.

Because the wheat producer-price series is available only quarterly, monthly flour and bread consumer prices were aggregated using arithmetic quarterly averages, yielding a balanced dataset of 44 observations with no missing values.

2.2. Sample overview and preliminary tests

The raw quarterly price series for wheat, flour, and bread were first transformed into natural logarithms to stabilize variance and interpret changes in proportional terms. Let (P_{it}) denote the observed price of commodity (i) in quarter (t), where ($i = 1, 2, 3$) refers respectively to wheat, flour, and bread. The logarithmic price series are defined as $(y_{it} = \ln P_{it})$, while the first differences, $(\Delta y = y_{it} - y_{it-1})$, represent quarterly log-price changes. Unit-root and stationarity test results are reported in Table 2.

The ADF and Phillips–Perron tests fail to reject the unit-root null for all three log-level price series, whereas both tests reject the null for the corresponding first differences at the 5% level. KPSS results reject trend stationarity for flour and bread in levels but not for wheat, indicating some ambiguity regarding the integration properties of the wheat price series. In first differences, the KPSS test does not reject stationarity for any of the three series. Taken together, the evidence is broadly consistent with treating the price series as $I(1)$; however, the wheat-level result and fixed-lag ADF sensitivity checks indicate finite-sample and specification uncertainty.

Cointegration among the price series was assessed using the Johansen trace test. Under conventional asymptotic critical values, the baseline specification rejects the null hypothesis of no cointegration ($r = 0$) but does not reject the null of at most one cointegrating vector ($r \leq 1$), implying a rank-one specification. Because the sample contains only 44 quarterly observations, the trace statistics were additionally evaluated using a small-sample adjustment. Under this adjustment, the null hypothesis of no cointegration is no longer rejected. The evidence for cointegration is therefore sample-sensitive, and the rank-one VECM is treated as a conditional baseline specification rather than as definitive

Table 2
Unit-root and stationarity tests for wheat, flour and bread price series.

Series	ADF level	ADF Δ	PP level	PP Δ	KPSS level	KPSS Δ
Wheat	−3.391	−4.102*	−2.271	−5.469*	0.114	0.073
Flour	−0.977	−4.074*	−1.114	−6.008*	0.159*	0.212
Bread	−1.511	−3.262*	−1.276	−6.500*	0.269*	0.376

Note: ADF and Phillips–Perron tests have a unit-root null hypothesis, whereas the KPSS test has a stationarity null hypothesis. An asterisk denotes rejection of the respective null hypothesis at the 5% level. ADF lag length was selected by AIC from a maximum of four lags.

evidence of a stable long-run relationship.

2.3. Model specification

The empirical framework is based on a VECM, which captures both the long-run equilibrium relationship among the price series and their short-run adjustment dynamics. Let

$$y_{1t} = \ln \text{wheat}_t, \quad y_{2t} = \ln \text{flour}_t, \quad y_{3t} = \ln \text{bread}_t,$$

where y_{1t} , y_{2t} and y_{3t} denote the natural logarithms of wheat, flour, and bread prices, respectively. These variables are collected in the vector

$$Y_t = \begin{pmatrix} y_{1t} \\ y_{2t} \\ y_{3t} \end{pmatrix}$$

The first-difference vector is defined as

$$\Delta Y_t = Y_t - Y_{t-1}$$

Lag-order criteria produced mixed results: the AIC and FPE selected three lags, whereas the HQ and Schwarz criteria selected one lag. Given the short quarterly sample and the need to preserve degrees of freedom, the final Johansen/VECM specification was estimated with two lags in levels, corresponding to one lag in first differences. Sensitivity was also examined under alternative deterministic specifications and lag lengths; these checks produced the same qualitative contrast between conventional and small-sample-adjusted rank inference. Conditional on the rank-one specification indicated by the conventional Johansen trace test, the baseline VECM is specified as

$$\Delta Y_t = \alpha \beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \varepsilon_t,$$

where $\beta' Y_{t-1}$ is the error-correction term (ECT), β is the cointegrating vector describing the long-run equilibrium relationship, α is the 3×1 loading vector measuring the speed and direction of adjustment toward equilibrium, Γ_1 is the 3×3 matrix of short-run coefficients, and ε_t is a vector of mean-zero innovations. Each element of α indicates how the corresponding price series responds to deviations from the long-run equilibrium, while each element of Γ_1 measures the effect of a lagged quarterly change in one price series on the current change in another. The baseline specification includes a linear trend restricted to the cointegrating relationship.

2.4. Statistical analysis

All analyses were performed in R. After accounting for lagged terms, the baseline dynamic equations contained 42 effective observations.

Dynamic responses were examined using Cholesky-orthogonalized impulse-response functions and forecast-error variance decompositions. The baseline recursive ordering was wheat $\rightarrow$ flour $\rightarrow$ bread. Because these results depend on the ordering of variables, alternative recursive orderings were examined as robustness checks. Generalized impulse-response functions and generalized forecast-error variance decompositions were also calculated as ordering-invariant descriptive robustness measures. Block-exogeneity tests were used to assess whether lagged changes in the other price variables jointly contributed to each equation. To assess whether a common nominal trend could account for part of the observed long-run association, unit-root and stationarity tests were also applied to the pairwise relative log-price series $\ln(\text{Flour}/\text{Wheat})$, $\ln(\text{Bread}/\text{Wheat})$, and $\ln(\text{Bread}/\text{Flour})$.

Because the asymmetric ECMs use the error-correction term derived from the maintained rank-one VECM, all subsequent asymmetry results are conditional on that rank-one specification. Asymmetric adjustment was examined using auxiliary parsimonious ECM specifications. For long-run asymmetry, the centred error-correction term was separated into positive and negative components, $ECT_{t-1}^+ = \max(ECT_{t-1}, 0)$ and $ECT_{t-1}^- = \min(ECT_{t-1}, 0)$, and equality of their coefficients was evalu-

ated using Wald tests.

For each price equation, the long-run asymmetric ECM is specified as

$$\Delta y_{j,t} = c_j + \lambda_j^+ ECT_{t-1}^+ + \lambda_j^- ECT_{t-1}^- + \Gamma_j' \Delta Y_{t-1} + u_{j,t},$$

where j denotes the wheat, flour or bread price equation. Long-run asymmetry is evaluated by testing $H_0: \lambda_j^+ = \lambda_j^-$.

Short-run asymmetry in the bread consumer-price equation was examined by separating lagged wheat and flour price changes into positive and negative components. The short-run asymmetric bread-price equation is specified as

$$\Delta bread_t = c + \lambda ECT_{t-1} + \theta_w^+ \Delta wheat_{t-1}^+ + \theta_w^- \Delta wheat_{t-1}^- + \theta_f^+ \Delta flour_{t-1}^+ + \theta_f^- \Delta flour_{t-1}^- + \phi \Delta bread_{t-1} + u_t.$$

Short-run asymmetry was evaluated by testing $H_0: \theta_w^+ = \theta_w^-$ and $H_0: \theta_f^+ = \theta_f^-$.

Both conventional and HC1 heteroskedasticity-robust Wald tests were calculated. These specifications are interpreted as auxiliary asymmetric ECMs rather than as formally estimated threshold-cointegration models.

Finally, an indicator equal to one from 2021 Q1 onward was included in the bread-price-change equation. Because this period encompasses several overlapping disturbances, including pandemic-related disruptions, the Ukraine-related commodity shock, and energy-price inflation, the indicator is interpreted as a broad post-2021 shock-period variable rather than as a COVID-19 dummy. Since it enters the first-difference equation, its coefficient represents a difference in average quarterly price growth rather than a one-time price-level shift. Structural stability was additionally evaluated using OLS-CUSUM tests and a BIC-based breakpoint procedure.

3. Results

3.1. Cointegration and VECM estimation

The conventional Johansen trace test for the baseline specification with a linear trend and two lags in levels supports a rank-one representation. The trace statistic for $r = 0$ was 47.757, exceeding the 5% critical value of 42.44, whereas the statistic for $r \leq 1$ was 23.690, below the corresponding critical value of 25.32. However, after small-sample adjustment, the $r = 0$ trace statistic decreased to 40.935 and no longer exceeded the 5% critical value (Table 3). The distinction between the conventional and adjusted results was also observed under alternative deterministic specifications and lag orders. Consequently, the following rank-one VECM results are interpreted conditionally on the conventional Johansen specification rather than as definitive evidence of cointegration.

The estimated cointegrating vector, normalized on the wheat-price series, was $1.000 \ln(\text{Wheat}) - 0.4873 \ln(\text{Flour}) - 2.4715 \ln(\text{Bread}) + 0.03184 \text{Trend}$. These coefficients describe the stationary linear combination implied by the conditional rank-one model and should not be interpreted as direct pass-through elasticities.

Table 3
Johansen trace test for the baseline specification.

Null hypothesis	Trace statistic	Small-sample-adjusted trace	5% critical value	Conventional decision	Adjusted decision
$r = 0$	47.757	40.935	42.44	Reject	Do not reject
$r \leq 1$	23.690	20.306	25.32	Do not reject	Do not reject
$r \leq 2$	4.812	4.125	12.25	Do not reject	Do not reject

Conditional on the rank-one specification, likelihood-ratio inference indicates substantial uncertainty in the normalized cointegrating coefficients. Exclusion of the bread coefficient is rejected ($\chi^2(1) = 5.050$, $p = 0.025$), whereas exclusion of the flour coefficient is not ($\chi^2(1) = 1.838$, $p = 0.175$). Exclusion of the restricted trend is borderline significant ($\chi^2(1) = 3.859$, $p = 0.049$), while the price-homogeneity restriction $\beta_W + \beta_F + \beta_B = 0$ is not rejected at the 5% level ($\chi^2(1) = 2.978$, $p = 0.084$).

Relative-price sensitivity checks yielded mixed evidence. ADF tests failed to reject a unit root for all three relative log-price series, whereas KPSS rejected stationarity for the flour-to-wheat and bread-to-flour relative prices but not for the bread-to-wheat relative price. Thus, the relative-price checks do not establish stationary relative prices and do not rule out a contribution of common inflationary trends to the observed nominal-price co-movement.

Fig. 2 plots the evolution of the ECT and the deviations from the estimated long-run relation over time.

Fig. 2 shows the centred error-correction term, calculated from the estimated long-run cointegrating relationship and expressed as deviations around its sample mean. Values close to zero indicate that the wheat–flour–bread price system is near its long-run equilibrium, whereas positive or negative values indicate temporary deviations from that equilibrium. Conditional on the rank-one specification, the centred ECT fluctuates around zero for most of the sample, with larger deviations occurring in the post-2021 period.

Table 4 presents the estimated short-run dynamics from the VECM equations for quarterly changes in wheat, flour, and bread prices.

The error-correction loading is statistically significant only in the flour consumer-price equation (0.4795, $p < 0.001$). Conditional on the rank-one VECM, this indicates that changes in the flour consumer-price series are most strongly associated with the correction of deviations from the estimated long-run relation. This finding cannot be attributed specifically to flour millers or to a mill-gate processing stage because the observed variable is a consumer price. The loading in the wheat producer-price equation is small and statistically insignificant, while the loading in the bread consumer-price equation is marginally significant at the 10% level.

In the short-run dynamics, lagged flour price changes have a negative and statistically significant effect on current flour price changes (-0.3235 , $p < 0.05$), indicating partial short-run correction within the flour-price segment. Other lagged differenced variables are not statistically significant at the 5% level. In the bread equation, the lagged own-price change is negative and marginally significant, suggesting some correction in bread consumer-price dynamics. Still, the baseline linear VECM does not provide strong evidence of direct short-run pass-through from lagged wheat or flour changes to bread prices. This issue is examined further through the asymmetric ECM tests reported below.

3.2. Model diagnostics

Fig. 3 presents the OLS-CUSUM empirical fluctuation processes for each of the three short-run equations in the VECM's companion VAR (wheat, flour, and bread).

In each panel, the horizontal axis is normalized time (from the first to the last observation), and the vertical axis shows the cumulative sum of recursive residuals. The dashed lines indicate the 95% confidence bounds: crossing them would signal a structural break or parameter instability. Across all three equations, the empirical fluctuation processes remain within the 95% confidence bounds. The corresponding OLS-CUSUM tests provide no evidence of parameter instability for the wheat ($p = 0.553$), flour ($p = 0.853$), or bread ($p = 0.665$) equations. In addition, the Portmanteau test for residual serial correlation up to lag 8 yielded $\chi^2(57) = 59.73$ ($p = 0.377$), providing no evidence of remaining residual autocorrelation. These diagnostics do not detect instability or residual serial dependence in the estimated short-run equations.

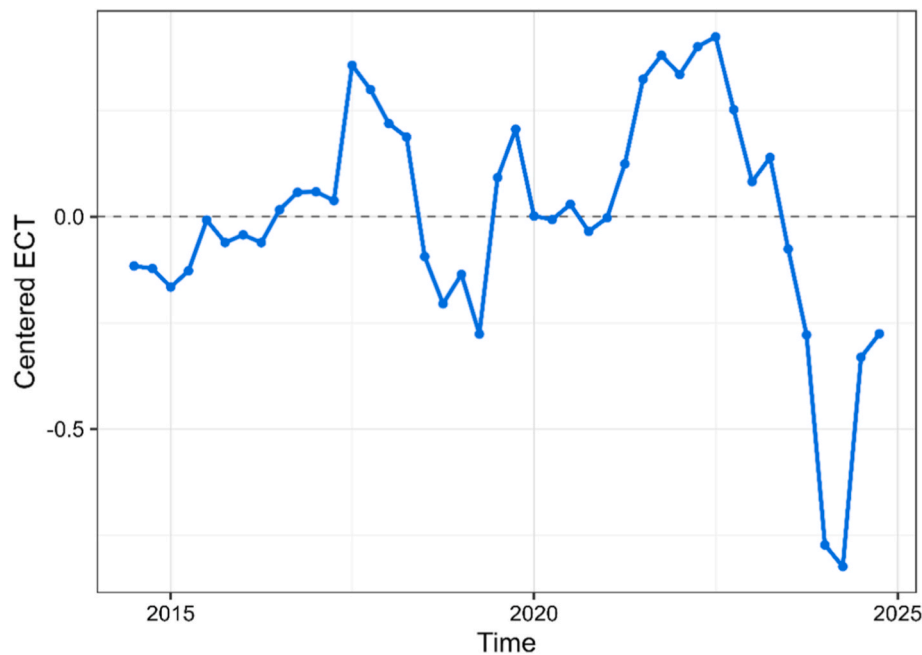


Fig. 2. Centred error-correction term from the conditional rank-one VECM specification.

Table 4

Parameter estimates from the conditional rank-one VECM.

Regressor	Δ Wheat (SE)	Δ Flour (SE)	Δ Bread (SE)
ECT_1 (long-run)	0.0172 (0.1008)	0.4795 (0.0997)***	0.0581 (0.0340).
Constant	0.0515 (0.2856)	1.3610 (0.2824)***	0.1780 (0.0964).
Δ Wheat _{t-1}	0.1375 (0.1640)	-0.0556 (0.1621)	0.0469 (0.0554).
Δ Flour _{t-1}	-0.0763 (0.1527)	-0.3235 (0.1510)*	-0.0199 (0.0516)
Δ Bread _{t-1}	-0.2413 (0.5786)	-0.8830 (0.5720)	-0.3401 (0.1954).

Note: N = 42 effective observations. Standard errors are shown in parentheses. The estimates are conditional on the rank-one VECM specification supported by conventional Johansen inference. $p < 0.10$; * $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$.

3.3. Dynamic transmission analysis

The dynamic responses of flour and bread prices to wheat price shocks were examined using orthogonalized impulse-response functions. The three-panel chart displays the dynamic responses of wheat, flour, and bread prices over a ten-quarter horizon following a one-standard-deviation innovation in wheat. Black lines trace the point-estimate IRFs, and red dashed lines represent 95% bootstrap confidence intervals. Time 0 on the x-axis corresponds to the shock period, and subsequent quarters show how each series evolves over the forecast horizon (Fig. 4).

A one-standard-deviation wheat-price innovation generates a positive and statistically supported own-price response throughout the ten-quarter horizon. Flour-price point estimates are initially slightly negative but turn positive from the second quarter onward, reaching 0.148 at horizon 10, while the bread-price response remains considerably smaller, reaching 0.028. However, the 95% confidence intervals for both downstream responses include zero at all horizons. The downstream IRF estimates should therefore be interpreted as qualitative dynamic patterns rather than statistically established responses.

The sensitivity analysis across alternative recursive orderings confirms that the bread response is not statistically significant at horizon 10 under any examined ordering. The flour response is also statistically insignificant under the baseline wheat–flour–bread ordering and most alternatives; it becomes only narrowly significant when flour is ordered before wheat, indicating some sensitivity to recursive identification.

Generalized impulse-response point estimates show a similar qualitative pattern, reaching 0.148 for flour and 0.028 for bread at horizon 10 following a wheat innovation. These generalized estimates are used descriptively because confidence intervals were not used for formal inference.

Block-exogeneity tests likewise provide no statistically significant evidence of short-run directional predictability. Lagged flour and bread changes are jointly insignificant in the wheat equation ($F(2,37) = 0.195$, $p = 0.824$), the lagged wheat change is insignificant in the flour equation ($F(1,37) = 0.118$, $p = 0.734$), and lagged wheat and flour changes are jointly insignificant in the bread equation ($F(2,37) = 0.454$, $p = 0.638$).

3.4. Forecast-error variance decomposition (FEVD)

Fig. 5 presents the forecast-error variance decomposition for wheat, flour, and bread prices over a ten-quarter horizon. The stacked bars show the proportion of forecast-error variance in each price series attributable to shocks in wheat, flour, and bread.

Forecast-error variance decomposition complements the impulse-response analysis by quantifying the relative contribution of each shock to forecast-error variance at different horizons. The results show that wheat price variability is almost entirely explained by its own shocks over the ten-quarter horizon. At horizon 10, own wheat shocks account for 99.73% of wheat forecast-error variance, while flour and bread shocks contribute only marginally. This is consistent with wheat prices being only weakly influenced by shocks to flour and bread within the estimated system.

For flour prices, own shocks dominate at very short horizons, explaining 99.85% of forecast-error variance at horizon 1. However, the contribution of wheat shocks increases steadily over time, reaching 45.47% by horizon 10. At the same horizon, flour's own shocks account for 21.43%, while bread shocks explain 33.10% of the variance. This suggests that wheat price shocks become increasingly important in explaining flour price variability over longer horizons.

For bread prices, own innovations remain the largest source of forecast-error variance throughout the horizon, declining from 97.09% at horizon 1 to 55.05% at horizon 10 under the baseline recursive ordering. Wheat innovations account for 37.48% of bread forecast-error

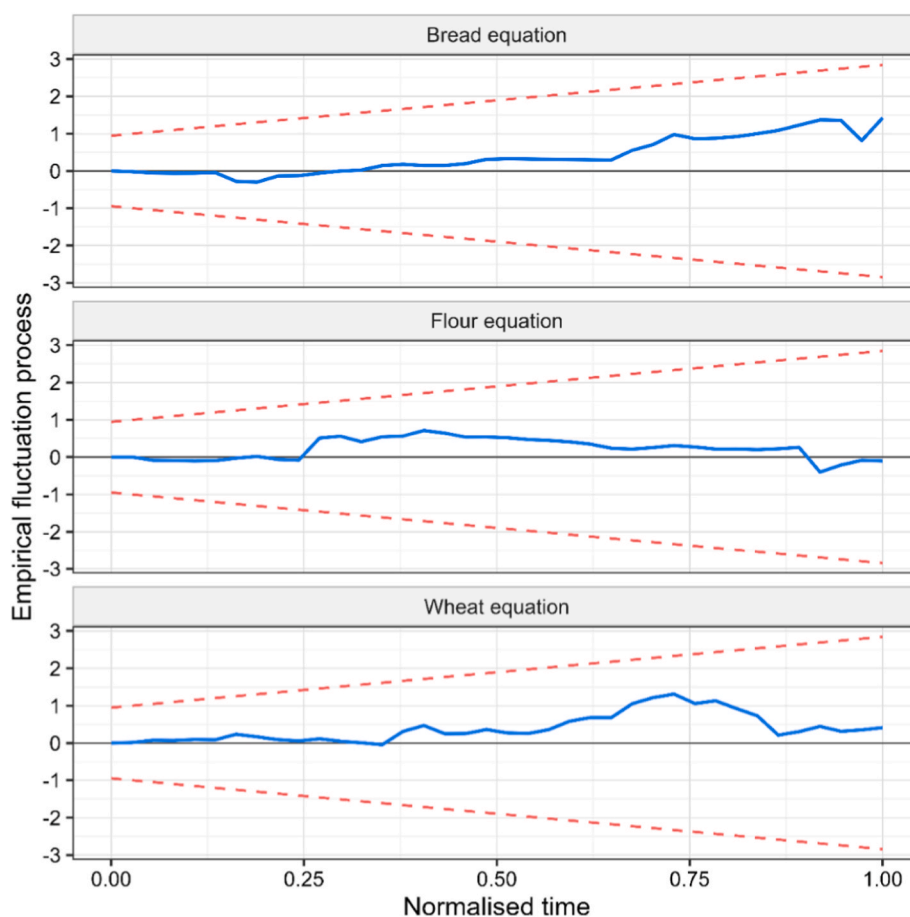


Fig. 3. OLS-CUSUM stability test for the companion VAR equations.

variance at horizon 10, while flour innovations account for 7.47%. Generalized FEVD yields a qualitatively similar result, with wheat innovations accounting for 32.82% of bread and 48.42% of flour forecast-error variance at horizon 10. These variance shares indicate increasing statistical importance of wheat innovations at longer forecast horizons. Still, they are descriptive decompositions rather than significance tests and therefore do not override the nonsignificant downstream IRF confidence intervals.

3.5. Asymmetric transmission and robustness checks

Conditional on the maintained rank-one specification, additional asymmetric ECM specifications were estimated to assess whether adjustment differs between positive and negative deviations or price changes. The centred error-correction term was decomposed into positive and negative deviations from the long-run equilibrium, and Wald tests were applied to determine whether these deviations are corrected at statistically different speeds. The results are shown in Table 5.

HC1-robust tests reject equality of positive and negative error-correction coefficients only in the bread consumer-price equation ($F(1,36) = 5.645$, $p = 0.023$); symmetry is not rejected for the wheat or flour equations. In the short-run bread-price equation, the formal equality test for positive and negative wheat-price changes does not reach the 5% significance level (HC1 $p = 0.058$), providing only marginal evidence of asymmetry at the 10% level. Positive and negative flour price changes do not differ significantly (HC1 $p = 0.426$).

The post-2021 indicator is positive and significant (0.0344, $p = 0.027$), indicating a higher average quarterly bread-price growth rate during the period beginning in 2021 Q1. Because this period includes several overlapping macroeconomic disturbances, the coefficient

should not be interpreted as an isolated COVID-19 effect or as a one-time level shift. The BIC-based procedure selects no permanent break in the bread-price equation.

4. Discussion

Bosnia and Herzegovina's wheat market is highly import-dependent: during 2014–2023, the average self-sufficiency ratio was 32.35%, while the average import-dependency ratio was 83.24%. This reliance on imports increases exposure to external commodity shocks, regional trade disruptions, and changes in input costs. Such vulnerability became particularly visible during the 2021–2022 shock period, when wheat prices rose sharply and flour prices exhibited even greater volatility, while bread price grew comparatively more slowly. These patterns are consistent with broader evidence that trade openness and import dependence shape cereal market behaviour and agricultural supply chain resilience. Zereyesus et al. [42] show that, in the case of wheat, trade openness may negatively affect cultivated area, yields, and total production, with the effects being especially relevant for countries with high import dependence. These findings are relevant to the Bosnia and Herzegovina context, where persistent production deficits and reliance on imports increase exposure to external price shocks and limit the capacity of domestic production to buffer volatility. Accordingly, the econometric results should be interpreted primarily as evidence concerning dynamic co-movement among the observed wheat producer-price and flour and bread consumer-price series within an import-dependent food system, rather than as direct identification of adjustment at specific processing or retail stages.

The long-run evidence requires a cautious interpretation. Under conventional asymptotic critical values, the Johansen trace test supports

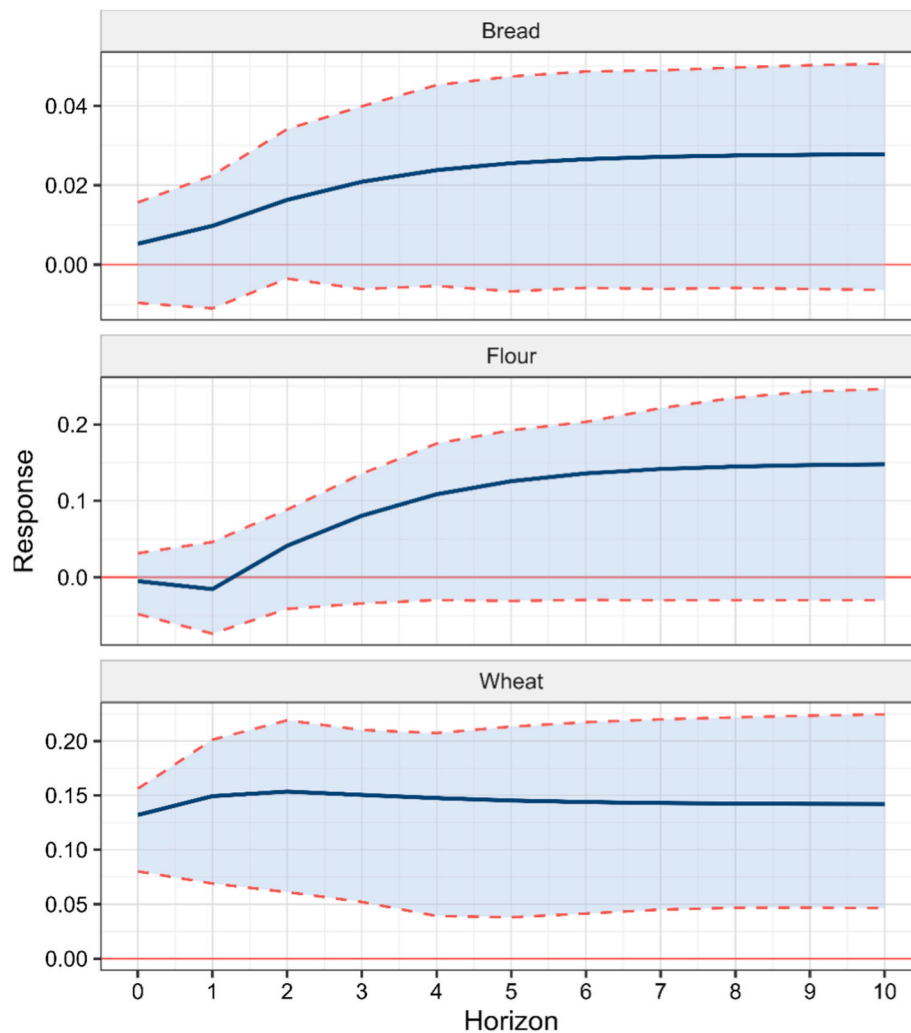


Fig. 4. Orthogonalized impulse-response functions to a one-standard-deviation wheat price shock.

one cointegrating relationship among wheat, flour, and bread prices. However, this conclusion is not preserved after small-sample adjustment, under which the null hypothesis of no cointegration is no longer rejected. The evidence for a stable long-run relationship should therefore be regarded as sample-sensitive rather than definitive, particularly given the limited number of quarterly observations. The wide and partly unbounded profile likelihood-ratio confidence sets for the normalized cointegrating coefficients further underline the limited precision of the estimated long-run relation in this small sample. Nevertheless, the conventional rank-one result can be considered in the context of evidence from other European wheat markets, where long-run price linkages have been documented despite episodes of substantial volatility. Bareith et al. [25], for example, report evidence of long-run cointegration between wheat producer prices and consumer prices of wheat flour in Hungary during the COVID-19 and Ukraine crisis periods, while Rumánková [27] finds that the Czech wheat–flour–bread chain remains linked in the long run even though adjustment is slow and partly asymmetric. In the case of Bosnia and Herzegovina, the results therefore suggest possible long-run co-movement, but the strength of this conclusion is constrained by small-sample uncertainty.

The impulse-response results reinforce the need to distinguish qualitative dynamic patterns from statistically supported effects. A one-standard-deviation wheat price innovation generates a statistically significant positive own response in wheat prices. The point estimates for flour and bread are consistent with weaker and delayed downstream responses: the flour response becomes positive after the first quarter. It

increases gradually over the forecast horizon, whereas the bread response remains considerably smaller. However, the 95% confidence intervals for both downstream responses include zero at all reported horizons. Sensitivity analysis across alternative recursive orderings likewise shows no statistically significant bread response at horizon 10, while the statistical support for the flour response depends on the ordering. Generalized impulse-response estimates preserve a similar qualitative pattern. Consequently, the evidence is consistent with delayed and attenuated downstream transmission but does not statistically establish that flour or bread prices respond significantly to wheat price innovations. A related pattern has been observed in Italy, where Ricci et al. [24] show that price transmission in wheat-based value chains is associated with downstream market structure, including the composition of the retail sector. The qualitative similarity should, however, be interpreted with caution because the present data do not separately identify processing costs, distribution margins, or retail price-setting mechanisms.

Conditional on the baseline rank-one VECM, the error-correction loading is statistically significant only in the flour consumer-price equation. This indicates that changes in the flour consumer-price series are most strongly associated with the correction of deviations from the estimated long-run relation. The result should not be interpreted as evidence that millers or the milling stage specifically restore equilibrium, because the observed flour variable is a consumer price that incorporates processing, packaging, distribution, wholesale, and retail components. By contrast, the wheat producer-price equation does not

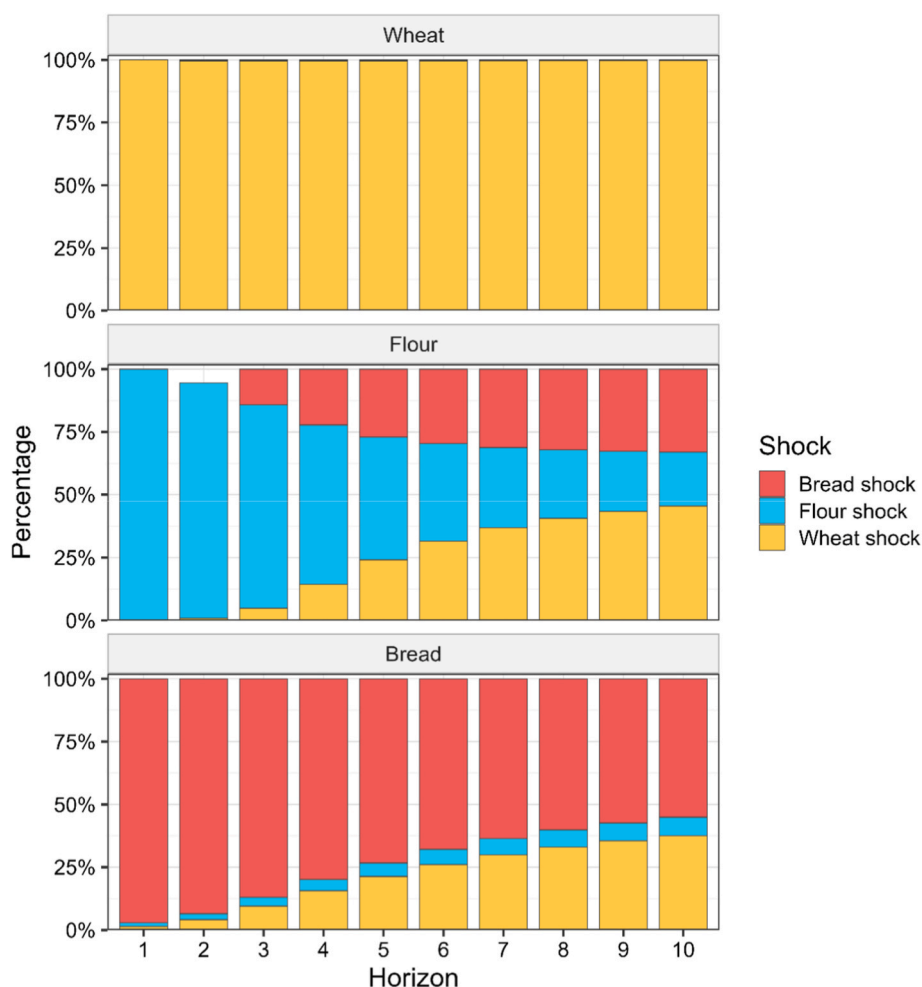


Fig. 5. Forecast-error variance decomposition for wheat, flour and bread prices.

exhibit statistically significant error-correction adjustment, while the bread consumer-price equation shows only marginal adjustment at the 10% level. The significant flour-price loading therefore identifies an adjustment in the observed consumer-price series rather than a specific firm-level or supply-chain mechanism.

The asymmetric ECM results provide the clearest evidence of non-uniform adjustment in the bread consumer-price equation. Under HC1-robust inference, equality of positive and negative error-correction coefficients is not rejected for the wheat or flour equations, whereas it is rejected for the bread equation. Conditional on the maintained rank-one specification, the HC1-robust Wald test indicates that bread consumer prices correct positive and negative deviations from the estimated long-run relation differently. This robustness refers to the covariance-based inference and does not remove the uncertainty surrounding the underlying cointegration rank. The short-run results are less conclusive. Although positive lagged wheat price changes have a statistically significant individual coefficient, the appropriate formal test directly compares the positive and negative wheat price coefficients. This equality test is not significant at the conventional 5% level and provides only marginal evidence of asymmetry at the 10% level. Therefore, the revised results do not support a strong conclusion that bread prices systematically respond more strongly to wheat-price increases than to wheat-price decreases in the short run. This distinction is important when relating the findings to the broader literature on asymmetric price transmission, in which market structure, adjustment costs, and other price-setting frictions have been discussed as possible explanations for asymmetric responses [14,15]. It also corresponds to empirical evidence from European wheat chains, where asymmetric or delayed downstream

adjustment has been documented in retail and bakery markets [24,27, 28]. However, the present results cannot identify market power, margins, or specific retail conduct from the available aggregate price series.

The FEVD results provide complementary descriptive evidence that wheat innovations become increasingly relevant for longer-horizon forecast uncertainty in the flour and bread consumer-price series. Generalized FEVD yields the same qualitative pattern. However, these variance decompositions are not significance tests and do not override the nonsignificant downstream IRF confidence intervals or the nonsignificant block-exogeneity tests.

In terms of the pre-specified hypotheses, H1 receives only conditional support: conventional Johansen inference favours a rank-one specification, but this result is not robust to small-sample adjustment. H2 receives qualitative rather than inferential support, because the flour and bread impulse-response point estimates are compatible with delayed and attenuated responses but do not differ significantly from zero. H3 is partially supported conditional on the maintained rank-one specification: HC1-robust inference indicates asymmetric equilibrium correction in the bread consumer-price equation, whereas short-run wheat-price asymmetry is only marginally supported at the 10% level.

From a policy perspective, the results support improved market monitoring rather than intervention targeted at a specific processing or retail stage. Monitoring should integrate domestic producer prices with import prices, exchange rates, major processing costs, and genuinely stage-specific flour and bread prices. Such information would help distinguish external cost transmission from other drivers of consumer-price adjustment.

Table 5

Asymmetric ECM estimates and robustness tests.

Panel A. Long-run asymmetric adjustment						
Equation	ECT ⁺ , estimate (SE)	p-value	ECT ⁻ , estimate (SE)	p-value	HC1 Wald <i>F</i> (1, 36)	HC1 p-value
Wheat	0.1851 (0.192)	0.342	-0.1009 (0.153)	0.514	1.110	0.299
Flour	0.1520 (0.182)	0.408	0.7097 (0.145)	<0.001	2.118	0.154
Bread	0.1798 (0.061)	0.006	-0.0274 (0.049)	0.578	5.645	0.023
Panel B. Short-run asymmetric bread-price equation						
Regressor	Estimate	SE	t-value	p-value		
Constant	0.0006	0.010	0.056	0.956		
<i>ECT</i> _{<i>t</i>-1}	0.0624	0.034	1.818	0.078		
$\Delta Wheat_{t-1}^{+}$	0.2052	0.101	2.040	0.049		
$\Delta Wheat_{t-1}^{-}$	-0.0966	0.091	-1.065	0.294		
$\Delta Flour_{t-1}^{+}$	0.0099	0.078	0.128	0.899		
$\Delta Flour_{t-1}^{-}$	-0.0408	0.066	-0.618	0.541		
$\Delta Bread_{t-1}$	-0.4706	0.204	-2.310	0.027		
HC1 Wald tests						
Null hypothesis	HC1 Wald <i>F</i>	df	p-value	Interpretation		
$\Delta Wheat_{t-1}^{+} = \Delta Wheat_{t-1}^{-}$	3.830	(1,35)	0.058	Marginal at 10%; not significant at 5%		
$\Delta Flour_{t-1}^{+} = \Delta Flour_{t-1}^{-}$	0.649	(1,35)	0.426	Not significant		
Panel C. Post-2021 shock-period robustness model						
Regressor	Estimate	SE	t-value	p-value		
Constant	0.0024	0.009	0.270	0.789		
<i>ECT</i> _{<i>t</i>-1}	0.0570	0.032	1.770	0.085		
$\Delta Wheat_{t-1}$	0.0416	0.052	0.794	0.432		
$\Delta Flour_{t-1}$	0.0002	0.049	0.003	0.997		
$\Delta Bread_{t-1}$	-0.4257	0.189	-2.259	0.030		
Post-2021 indicator	0.0344	0.015	2.313	0.027		

Note: N = 42 for all models. Wald tests are based on HC1 heteroskedasticity-robust covariance estimates. All asymmetric ECM results are conditional on the maintained rank-one cointegrating specification. *HC1-robust* refers to heteroskedasticity-robust covariance inference and does not imply robustness of the underlying cointegration rank.

5. Conclusions

This study examined dynamic relationships between wheat producer prices and consumer prices of flour and bread in Bosnia and Herzegovina using 44 quarterly observations from 2014 Q1 to 2024 Q4. Conventional Johansen inference supports a rank-one specification, but this finding is not robust to small-sample adjustment and should therefore be regarded as suggestive. Conditional on the rank-one VECM, the flour consumer-price equation exhibits the strongest error-correction loading. Impulse-response point estimates are compatible with delayed and attenuated downstream responses to wheat-price innovations, but the flour and bread responses are not statistically significant under the baseline ordering. Conditional on the maintained rank-one specification, HC1-robust inference indicates asymmetric equilibrium correction in the bread consumer-price equation. In contrast, short-run wheat-price asymmetry is only marginally supported at the 10% level.

The findings therefore support cautious interpretation of dynamic price co-movement rather than stage-specific conclusions about processing, retail conduct, or market power. For policy purposes, the principal priority is improved market information that combines domestic producer prices with import prices, exchange rates, major processing costs, and genuinely stage-specific flour and bread prices. Such data, together with longer and higher-frequency time series, would provide a stronger basis for assessing price transmission and designing targeted market interventions.

6. Limitations

Several limitations should be considered when interpreting the results. The sample contains only 44 quarterly observations, which limits statistical power, and aggregating monthly flour and bread consumer prices to quarterly averages may attenuate short-run responses.

Moreover, the flour and bread variables are consumer prices that incorporate multiple processing and distribution components and therefore do not identify specific supply-chain stages or firm-level margins. The model also omits potentially relevant common drivers, including import prices, exchange rates, energy and labour costs, quantities, and market margins. Common inflationary trends may also contribute to the observed co-movement in nominal prices, as the relative-price sensitivity checks provide mixed evidence. Orthogonalized IRFs and FEVDs depend on recursive identification; however, alternative orderings and generalized measures were examined, while the asymmetric ECMs represent parsimonious auxiliary specifications rather than formally estimated threshold-cointegration models. The short sample also limits the use and power of fully break-robust unit-root and cointegration procedures, despite the structural-stability diagnostics reported here. Longer, higher-frequency, and genuinely stage-specific datasets would allow future research to apply more flexible nonlinear models.

CRediT authorship contribution statement

Vesna Mrdalj: Writing – review & editing, Writing – original draft, Investigation, Formal analysis, Conceptualization. **Đorđe Stojisavljević:** Writing – review & editing, Writing – original draft, Methodology, Data curation, Conceptualization. **Marija Gjosheva Kovachevikj:** Writing – review & editing, Writing – original draft, Supervision, Data curation. **Aleksandar Ostojić:** Writing – review & editing, Writing – original draft, Visualization, Supervision, Data curation. **Tamás Mizik:** Writing – review & editing, Writing – original draft, Supervision, Resources, Funding acquisition, Formal analysis.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

Data will be made available on request.

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