**Merve KOCAMAN, PhD**
mervealtin@anadolu.edu.tr
Anadolu University, Eskisehir, Türkiye

# Do Wages Chase Prices or Prices Chase Wages? Revisiting the Wage–Price Spiral in Türkiye with QARDL Approach

**Abstract.** *This study examines the validity of the wage–price spiral in the Turkish economy using quarterly data from 1988: Q1 to 2025: Q1 within a Quantile Autoregressive Distributed Lag (QARDL) framework, complemented by the Fourier Toda–Yamamoto (F-TY) causality test. The analysis incorporates oil prices, the exchange rate, unemployment, and real GDP as control variables to capture broader macroeconomic dynamics. The results show that wage increases exert a consistent upward pressure on inflation across the entire distribution, although the magnitude of this effect varies depending on the prevailing inflation regime. At the same time, inflation pass-through to wages is observed across all quantiles, with the effect particularly strong in the upper quantiles, reflecting stronger bargaining power and indexation mechanisms. The F-TY causality test further confirms statistically significant bidirectional causality between wages and inflation, providing robust evidence of a wage–price spiral in Türkiye. These findings highlight the risk that persistent inflation can fuel a self-reinforcing cycle of rising wages and prices, underscoring the need for credible disinflation strategies and coordinated policy measures to safeguard macroeconomic stability.*

**Keywords**: *wage-price spiral, QARDL, Fourier causality.*

**JEL Classification:** E31, J30, C32.

| **Received: 29 September 2025** | **Revised: 2 June 2026** | **Accepted: 7 June 2026** |
| --- | --- | --- |

## 1. Introduction

The wage–price spiral is an important problem for macroeconomic stability especially in developing countries. In environments with high inflation, weak labour markets, and limited institutional strength, the interaction between wages and prices can harm both sustainable economic growth and fair income distribution. When workers demand higher wages to protect their real purchasing power and firms pass these higher costs on to prices, inflationary pressures grow stronger, creating a vicious cycle that disrupts economic balance. As stated by Zeira (1989), this spiral generates inflationary inertia, as wage increases trigger price hikes, which subsequently drive further wage demands, thereby allowing current inflation to spill over into future inflation rates.

The underlying mechanisms of the wage–price spiral can be broadly classified into two principal sources. First one is the demand-pull mechanism. An increase in aggregate demand was thought to raise output and employment, prompting firms to

DOI: 10.24818/18423264/60.2.26.03

seek higher prices and workers to demand higher wages. This process could initiate a wage–price spiral, which would persist until demand-pull inflation reduced real money balances sufficiently to restore the economy to its steady state. Secondly, the spiral could arise from workers' efforts to raise their real wages, firms' attempts to expand profit margins, or from both parties striving to maintain the existing wage and price levels in the face of an adverse supply shock. Such dynamics would likewise trigger a wage–price spiral, generate "cost-push" inflation, and through the erosion of real money balances ultimately lead to a recession (Blanchard, 1986: 543).

This study investigates the validity of the wage–price spiral in the context of the Turkish economy. To this end, the Quantile Autoregressive Distributed Lag (QARDL) model is employed to assess the impact of wage increases on inflation under different inflation regimes, as well as to explore the effects of inflationary pressures across various wage values. The analysis is based on quarterly data spanning the period 1988: Q1–2025: Q1, comprising 149 observations. In addition, to investigate the causal relationships among the variables, the Fourier Toda–Yamamoto causality test, which accounts for structural breaks, is applied.

The remainder of the study is organised as follows: Section 2 presents a short review of the empirical literature; Section 3 introduces the data and methodology; Section 4 reports the estimation results; and finally, Section 5 provides the concluding remarks.

## 2. Empirical Literature

Extensive empirical research has been conducted on price wage spiral, yet the evidence remains mixed, reflecting differences in methodological choices, sample periods, and country-specific dynamics. Some of the key contributions from this literature are summarised below.

Hoxha (2010) investigated the causal relationship between prices and wages in the EU-27 using OLS and VECM methods. The results offer strong evidence of a bidirectional causal link between prices and wages, both in the long run and the short run. Saunders and Denniss (2022) analyse Australian data to assess whether wage increases drive inflation. They find that, since wages account for only a limited share of business costs (on average 25.3%), a 5% wage increase would raise prices by merely 1.3–1.9%. Thus, there is no strong evidence that wage increases trigger a wage–price spiral or significantly accelerate inflation. Alvarez et al. (2024) investigated the wage–price spiral in 38 advanced economies using quarterly data from 1960:Q1 to 2021:Q4. Their models incorporated the consumer price index (CPI), nominal wages, unemployment rate, energy price index, productivity, and real GDP. Applying a panel fixed effects framework, they concluded that an acceleration in nominal wages amid rising inflation does not automatically imply the onset of a persistent wage–price spiral. Bernanke and Blanchard (2025) addresses question of U.S. pandemic-era inflation by developing and estimating a dynamic model of prices, wages, and both short- and long-term inflation expectations. The results indicate that, contrary to initial fears of an overheated labour market fuelling inflation, the sharp rise in inflation beginning in 2021 was largely driven by price shocks independent of wages.

Empirical research focusing on Türkiye likewise reveals varying outcomes. Başkaya and Özmen (2013) analysed the period 2003:Q1–2012:Q2 using regression techniques and found that increases in real minimum wages exert upward pressure on inflation. Abdioğlu (2015) investigates the dynamics of nominal wage and price adjustments in Türkiye in the face of aggregate demand shocks, utilising quarterly data spanning 1998–2012. Applying a stepwise regression framework that incorporates nominal GDP, public expenditures, M1 money supply, oil prices, nominal and real wage indices, the manufacturing production index, CPI, and interest rates, the study found that because nominal adjustments in labour and goods markets do not occur simultaneously, the wage–price spiral unfolds only gradually. Using monthly data spanning January 2005 to March 2017, Biçerli and Kocaman (2019) examined the impact of minimum wage increases on prices, unemployment, and economic growth. They applied the ARDL model and ECM-based Granger causality analysis, and found that increases in minimum wages lead to higher inflation and unemployment, with the causality results further reinforcing these findings. Akgül and Bükey (2020) employed ARDL bounds testing and the Toda–Yamamoto causality approach for the period 1987–2018, and their results revealed a bidirectional causal relationship between gross minimum wages and inflation. Özer and Gülşen (2025) investigated the relationship between minimum wages and inflation in Türkiye over the period 1969–2022 by employing the Fourier Toda–Yamamoto causality test. The findings indicate a bidirectional relationship between the inflation rate and nominal minimum wages, and a unidirectional relationship from real minimum wages to inflation.

Unlike previous studies, this research covers a broad time span from 1988:Q1 to 2025:Q1, comprising 149 observations. It is considered that the impact of wage increases may vary across low, medium, and high inflation regimes, and, similarly, that the effects of price increases may differ across low, middle, and high income values. Using the QARDL approach, the study aims to estimate the asymmetric dynamic relationships among the variables and thus contribute to the existing literature.

## 3. Data and Methodology

### *3.1 Data*

This study utilises Turkish quarterly data spanning the period 1988Q1–2025Q1 to examine the price–wage spiral model. The variables employed in the analysis are the consumer price index and the gross wages–salaries index. In addition, oil prices, real GDP, exchange rate, and unemployment rate are included in the model as control variables. All the data is obtained from DataStream except gross wages-salaries index. This index was obtained from the TURKSTAT database. The TRAMO/SEATS seasonal adjustment is applied to the series that exhibit seasonality.

Than to mitigate potential issues of heteroscedasticity and volatility, all variables are expressed in their natural logarithmic form. The empirical analyses were conducted using GAUSS 24 and EViews 14 software packages.

Accordingly, Equations (1) and (2) set out the proposed model, which explores the relationship between the dependent and independent variables.

$$linf = f(lwsi, loil, lunemp, lexc, lgdp) \tag{1}$$

$$\mathrm{lwsi} = \mathrm{f(lcpi, loil, lunemp, lexc, lgdp)} \tag{2}$$

### *3.2 Methodology*

To investigate the dynamics of the wage–price spiral, the Quantile Autoregressive Distributed Lag (QARDL) model, developed by Cho et al. (2015), is employed. This method provides an opportunity to observe the effects of wage increases across the low, medium and high inflation regimes in our model.
The empirical investigation commenced with an evaluation of the time-series properties of the variables, specifically assessing the presence of unit roots. Stationarity was formally tested through the Augmented Dickey-Fuller (ADF) test (Dickey & Fuller, 1981) and the Phillips-Perron (PP) test (Perron, 1989). In instances where variables were found to be integrated of order one – non-stationary in levels but stationary after first differencing – cointegration techniques were subsequently employed to ascertain the existence of long-run equilibrium relationships among the variables.

As a starting point, the traditional ARDL models are formulated as presented in Equations 3 and 4:

$$linf_t = c_0 + \sum_{i=1}^{p1} \mu_{1i}\,\Delta linf_{t-i} + \sum_{i=0}^{p2} \mu_{2i}\,\Delta lwsi_{t-i} + \sum_{i=0}^{p3} \mu_{3i}\,\Delta loil_{t-i} + \sum_{i=0}^{p4} \mu_{4i}\,\Delta lgdp_{t-i} + \sum_{i=0}^{p5} \mu_{5i}\,\Delta lunemp_{t-i} + \sum_{i=0}^{p6} \mu_{6i}\,\Delta lexc_{t-i} + u_{1t} \tag{3}$$

$$lwsi_t = c_1 + \sum_{i=1}^{q1} \eta_{1i}\,\Delta lwsi_{t-i} + \sum_{i=0}^{q2} \eta_{2i}\,\Delta linf_{t-i} + \sum_{i=0}^{q3} \eta_{3i}\,\Delta loil_{t-i} + \sum_{i=0}^{q4} \eta_{4i}\,\Delta lgdp_{t-i} + \sum_{i=0}^{q5} \eta_{5i}\,\Delta lunemp_{t-i} + \sum_{i=0}^{q6} \eta_{6i}\,\Delta lexc_{t-i} + u_{2t} \tag{4}$$

where $u_t{}'s$ denote the white noise disturbance terms, characterised by the minimal set ( $linf_t$ , $lwsi_t$ , $loil_t$, $lgdp_t$, $lunemp_t$, $lexc_t$, $linf_{t-1}$, $lwsi_{t-1}$, $loil_{t-1}$, $lgdp_{t-1}$, $lunemp_{t-1}$) wh ile $p1\ldots p6$ and $q1\ldots q6$ correspond to the lag orders selected according to the Akaike Information Criterion. Equations (3) and (4) are extended and reformulated by Cho et al. (2015) to derive the basic specification of the QARDL model, which is presented as follows:

$$Q_{linf_t} = c_0(\tau) + \sum_{i=1}^{n1} \varphi_1(\tau)\,\Delta linf_{t-i} + \sum_{i=0}^{n2} \varphi_2(\tau)\,\Delta lwsi_{t-i} + \sum_{i=0}^{n3} \varphi_3(\tau)\,\Delta loil_{t-i} + \sum_{i=0}^{n4} \varphi_4(\tau)\,\Delta lgdp_{t-i} + \sum_{i=0}^{n5} \varphi_5(\tau)\,\Delta lunemp_{t-i} + \sum_{i=0}^{n6} \varphi_6(\tau)\,\Delta lexc_{t-i} + \varepsilon_{1t}(\tau) \quad (5)$$

$$Q_{lwsi_t} = c_1(\tau) + \sum_{i=1}^{m1} \gamma_1(\tau)\,\Delta lwsi_{t-i} + \sum_{i=0}^{m2} \gamma_2(\tau)\,\Delta linf_{t-i} + \sum_{i=0}^{m3} \gamma_3(\tau)\,\Delta loil_{t-i} + \sum_{i=0}^{m4} \gamma_4(\tau)\,\Delta lgdp_{t-i} + \sum_{i=0}^{m5} \gamma_5(\tau)\,\Delta lunemp_{t-i} + \sum_{i=0}^{m6} \gamma_6(\tau)\,\Delta lexc_{t-i} + \varepsilon_{2t}(\tau) \quad (6)$$

Here $\varepsilon_{1t}(\tau)$ equals $linf_t - Q_{linf_t}[\tau/F_{t-1}]$ and $Q_{linf_t}(\tau/F_{t-1})$ refers to the $\tau_{th}$ quantile of $lcpi_t$ conditional on the information set $F_{t-1}$ and 10< τ <1 denotes the quantile. Similarly, $\varepsilon_{2t}(\tau)$ equals $lwsi_t - Q_{lwsi_t}[\tau/F_{t-1}]$ and $Q_{lwsi_t}(\tau/F_{t-1})$ refers to the $\tau_{th}$ quantile of $lwsi_t$ conditional on the information set $F_{t-1}$. This study conducts estimations across nine successive quantiles of the conditional distribution, specifically at τ = 0.10, 0.20, 0.30, 0.40, 0.50, 0.60, 0.70, 0.80, and 0.90. Due to the potential presence of autocorrelation in the error term, equations (5) and (6) can be re-specified as follows for the QARDL analysis.

$$Q_{\Delta linf_t} = c_0(\tau) + \rho linf_{t-i} + \omega_{lwsi} lwsi_{t-1} + \omega_{loil} loil_{t-1} + \omega_{lgdp} lgdp_{t-1} + \omega_{lunemp} lunemp_{t-1} + \omega_{lexc} lexc_{t-1} + \sum_{i=1}^{y1} \vartheta_1\,\Delta linf_{t-i} + \sum_{i=0}^{y2} \vartheta_2\,\Delta lwsi_{t-i} + \sum_{i=0}^{y3} \vartheta_3\,\Delta loil_{t-i} + \sum_{i=0}^{y4} \vartheta_4\,\Delta lgdp_{t-i} + \sum_{i=0}^{y5} \vartheta_5\,\Delta lunemp_{t-i} + \sum_{i=0}^{y6} \vartheta_6\,\Delta lexc_{t-i} + \epsilon_{1t}(\tau) \quad (7)$$

$$Q_{\Delta lwsi_t} = c_1(\tau) + \rho lwsi_{t-i} + \emptyset_{lcpi} linf_{t-1} + \emptyset_{loil} loil_{t-1} + \emptyset_{lgdp} lgdp_{t-1} + \emptyset_{lunemp} lunemp_{t-1} + \emptyset_{lexc} lexc_{t-1} + \sum_{i=1}^{z1} \delta_1(\tau)\,\Delta linf_{t-i} + \sum_{i=0}^{z2} \delta_2(\tau)\,\Delta lwsi_{t-i} + \sum_{i=0}^{z3} \delta_3(\tau)\,\Delta loil_{t-i} + \sum_{i=0}^{z4} \delta_4(\tau)\,\Delta lgdp_{t-i} + \sum_{i=0}^{z5} \delta_5(\tau)\,\Delta lunemp_{t-i} + \sum_{i=0}^{z6} \delta_6(\tau)\,\Delta lexc_{t-i} + \epsilon_{2t}(\tau) \quad (8)$$

In order to derive the QARDL error correction model (QARDL-ECM), equations (7) and (8) must be reformulated as presented in equation (9) and (10) below:

$$
\begin{aligned}
Q_{\Delta linf_t} = {} & c_0(\tau) + \rho(\tau)(linf_{t-i} + \beta_{lwsi}(\tau)lwsi_{t-1} + \beta_{loil}(\tau)loil_{t-1} + \beta_{lgdp}(\tau)lgdp_{t-1} \\
& + \beta_{lunemp}(\tau)lunemp_{t-1} + \beta_{lexc}(\tau)lexc_{t-1}) + \sum_{i=1}^{y1} \alpha(\tau)\, \Delta linf_{t-i} \\
& + \sum_{i=0}^{y2} \gamma(\tau)\, \Delta lwsi_{t-i} + \sum_{i=0}^{y3} \delta(\tau)\, \Delta loil_{t-i} + \sum_{i=0}^{y4} \vartheta(\tau)\, \Delta lgdp_{t-i} \\
& + \sum_{i=0}^{y5} \mu(\tau)\, \Delta lunemp_{t-i} + \sum_{i=0}^{y6} \omega(\tau)\, \Delta lexc_{t-i} + \epsilon_{1t}(\tau) \qquad (9)
\end{aligned}
$$

$$
\begin{aligned}
Q_{\Delta lwsi_t} = {} & c_1(\tau) + \rho(\tau)(lwsi_{t-i} + \beta_{lwsi}(\tau)linf_{t-1} + \beta_{loil}(\tau)loil_{t-1} + \beta_{lgdp}(\tau)lgdp_{t-1} \\
& + \beta_{lunemp}(\tau)lunemp_{t-1} + \beta_{lexc}(\tau)lexc_{t-1}) + \sum_{i=1}^{z1} \varphi(\tau)\, \Delta lwsi_{t-i} \\
& + \sum_{i=0}^{z2} \sigma(\tau)\, \Delta linf_{t-i} + \sum_{i=0}^{z3} \partial(\tau)\, \Delta loil_{t-i} + \sum_{i=0}^{z4} \eta(\tau)\, \Delta lgdp_{t-i} \\
& + \sum_{i=0}^{z5} \zeta(\tau)\, \Delta lunemp_{t-i} + \sum_{i=0}^{z6} \psi(\tau)\, \Delta lexc_{t-i} \\
& + \epsilon_{2t}(\tau) \qquad (10)
\end{aligned}
$$

Using the delta method, the cumulative short-run effect of past inflation on current inflation is represented by $\alpha^* = \sum_{i=1}^{y1} \alpha_i$ , for the first model. Similarly, the cumulative short-run effects of current and lagged values of wages, oil prices, gross domestic product, unemployment, and the exchange rate on the current level of inflation are denoted by $\gamma^* = \sum_{i=1}^{y2} \gamma_i$ , $\delta^* = \sum_{i=1}^{y3} \delta_i$ , $\vartheta^* = \sum_{i=1}^{y4} \vartheta_i$ , $\mu^* = \sum_{i=1}^{y5} \mu_i$ , $\omega^* = \sum_{i=1}^{y6} \omega_i$ , respectively. In addition, the long-run effects of wages, oil prices, GDP, unemployment, and the exchange rate on inflation are measured as follows: $\beta_{lwsi*} = -\frac{\beta_{lwsi}}{\rho}$, $\beta_{loil*} = -\frac{\beta_{loil}}{\rho}$, $\beta_{lgdp*} = -\frac{\beta_{lgdp}}{\rho}$, $\beta_{lunemp*} = -\frac{\beta_{lunemp}}{\rho}$, $\beta_{lexc*} = -\frac{\beta_{lexc}}{\rho}$. The error correction term (ρ) should be negative and statistically significant.

To assess the asymmetric effects of wages, oil prices, GDP, unemployment, and the exchange rate on inflation across different quantiles, the Wald test, based on the chi-squared distribution, is applied, capturing both short-run and long-run dynamics. For example, the following null hypothesis was tested for the speed of the adjustment parameter (ρ*):

$H_0: \rho * (0.10) = \rho * (0.20) = \rho * (0.30) = \rho * (0.40) = \rho * (0.50) = \rho * (0.60) = \rho * (0.70) = \rho * (0.80) = \rho * (0.90)$

The same hypothesis was also tested for the long-run parameters ($\beta_{lwsi*}, \beta_{loil*}, \beta_{lgdp*}, \beta_{lunemp*}, \beta_{lexc*}$ ) as well as for the short-run parameters. The same procedures have also been applied to the second model.

*3.2.1 Causality test*

In the second stage of the analysis, the causal relationships are examined. In line with the Toda–Yamamoto approach, this method allows for the analysis of non-stationary time series in levels, thereby avoiding the loss of long-run information. Accordingly, the Fourier Toda–Yamamoto (F-TY) causality test is employed to investigate the causal relationships among the variables. The TY test extends the VAR model by including an additional lag equal to the highest order of integration among the series. Thus, the standard VAR model in the F-TY framework can be expressed as VAR(p + d), where *p* denotes the optimal lag length, and *d* represents the maximum integration order. The model proposed by Nazlıoğlu et al. (2016), incorporating a single frequency component, is specified as follows:

$$y_t = \alpha(0) + \gamma_1 \sin\left(\frac{2\pi kt}{T}\right) + \gamma_2 \cos\left(\frac{2\pi kt}{T}\right) + \beta_1 \gamma_{t-1} + \cdots + \beta_{p+d} \gamma_{t-(p+d)} + \epsilon_t \qquad (11)$$

Equation (5) presents an important specification challenge, namely, the selection of the optimal number of Fourier frequencies and the appropriate lag length. To resolve this, the Akaike Information Criterion (AIC) is employed to determine the optimal values of *p* (lag length) and *k* (number of Fourier frequencies) for the model. Although the Wald statistic is traditionally used to test the null hypothesis in this framework, Lütkepohl (2005) recommends the use of the F-statistic, particularly in small samples. This preference arises because the chi-squared distribution underlying the Wald test often provides a poor approximation of the finite-sample distribution of the causality test statistic. By contrast, the F-distribution, with its heavier tails, yields a more reliable approximation under such conditions (Nazlıoğlu et al., 2016: 172).

To improve the reliability and finite-sample properties of the inference, the bootstrap distribution of the F-statistic is generated using a residual-based bootstrap procedure. The null hypothesis tested assumes no causal relationship between the variables. The final conclusion is then determined by whether this null hypothesis is rejected, based on the critical values obtained from the bootstrap simulation.

## 4. Estimation Results

The results of the ADF and PP unit root tests are presented in Table 1. The results of the ADF and PP unit root tests indicate that the majority of the variables are non-stationary in their level form but become statistically significantly stationary after first differencing. In both the constant and constant & trend specifications, stationarity at levels is limited and not consistently supported across the two tests, suggesting that most series are integrated of order one, I(1).

**Table 1. Unit root test results**

| Variables | ADF | | PP | |
|---|---|---|---|---|
| | Constant (prob.) | Constant and Trend (prob.) | Constant (prob.) | Constant and Trend (prob.) |
| linf | -1.65 (0.45) | -2.42 (0.36) | -3.59 (0.00) | -2.18 (0.49) |
| Δlinf | -2.92 (0.04)* | -5.4898 (0.00)* | - | -6.15 (0.00)* |
| lwsi | -1.48 (0.53) | -2.59 (0.28) | -2.77 (0.06) | -2.50 (0.32) |
| Δlwsi | -3.40 (0.01)* | -3.76 (0.02)* | -9.03 (0.00)* | -10.27 (0.00)* |
| lexc | -2.72 (0.07) | -1.94 (0.62) | -2.86 (0.05) | -2.00 (0.59) |
| Δlexc | -8.04 (0.00)* | -8.40 (0.00)* | -8.15 (0.00)* | -8.47 (0.00)* |
| lunemp | -2.05 (0.26) | -2.56 (0.29) | -1.91 (0.32) | -2.15 (0.50) |
| Δlunemp | -4.46 (0.00)* | -4.48 (0.00)* | -4.57 (0.00)* | -4.55 (0.00)* |
| loil | -1.71 (0.42) | -2.69 (0.24)** | -1.55 (0.50) | -2.42 (0.36) |
| Δloil | -10.20 (0.00)* | -10.17 (0.00)* | -10.10 (0.00)* | -10.07 (0.00)* |
| lgdp | 0.21 (0.97) | -3.25 (0.07) | 0.46 (0.98) | -3.32 (0.06) |
| Δlgdp | -13.34 (0.00)* | -13.32 (0.00)* | -13.67 (0.00)* | -13.73 (0.00)* |

*Source*: Author’s own calculation.

These findings point to a data structure that is suitable for cointegration analysis and for applying models such as QARDL, which can accommodate a mix of I(0) and I(1) variables. Table 2 reports the outcomes of the cointegration bounds test, where the null hypothesis of no cointegration is examined, and subsequently presents the QARDL model results for Model 1.

**Table 2. Quantile ARDL results for model 1**

**Panel A: Bounds test statistics**

| 1% | | 5% | | 10% | | |
|---|---|---|---|---|---|---|
| I(0) | (I1) | (I0) | I(1) | I(0) | (I1) | F-stat: 28.644 |
| 2.39 | 3.38 | 2.08 | 3 | 2.39 | 3.38 | |

**Panel B: Quantile ARDL**

| | ***Quantiles*** | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Variables | q0.10 | q0.20 | q0.30 | q0.40 | q0.50 | q0.60 | q0.70 | q0.80 | q0.90 |
| $\boldsymbol{c_0(\tau)}$ | 4.33[c] (2.29) | 4.89[a] (1.81) | 7.41[a] (1.55) | 7.15[a] (1.56) | 8.61[a] (1.51) | 8.74[a] (1.43) | 9.40[a] (1.16) | 10.06[a] (1.14) | 10.45[a] (1.08) |
| $\boldsymbol{\rho(\tau)}$ | -0.08[a] | -0.10[a] | -0.11[a] | -0.11[a] | -0.11[a] | -0.12[a] | -0.15[a] | -0.15[a] | -0.16[a] |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | (0.01) | (0.01) | (0.00) | (0.00) | (0.00) | (0.00) | (0.0) | (0.01) | (0.01) |
| $\boldsymbol{\beta_{lwsi*}(\tau)}$ | 0.42[a] (0.13) | 0.50[a] (0.10) | 0.55[a] (0.09) | 0.43[a] (0.09) | 0.39[a] (0.09) | 0.44[a] (0.08) | 0.54[a] (0.07) | 0.51[a] (0.06) | 0.31[a] (0.06) |
| $\boldsymbol{\beta_{loil*}(\tau)}$ | 0.03 (0.08) | 0.08 (0.06) | 0.07 (0.05) | 0.10[c] (0.05) | 0.15[a] (0.05) | 0.17[a] (0.04) | 0.20[a] (0.03) | 0.21[a] (0.03) | 0.28[a] (0.03) |
| $\boldsymbol{\beta_{lgdp*}(\tau)}$ | -0.20 (0.23) | -0.30 (0.10) | -0.51[a] (0.15) | -0.42[a] (0.16) | -0.52[a] (0.15) | -0.55[a] (0.14) | -0.69[a] (0.11) | -0.73[a] (0.11) | -0.80[a] (0.11) |
| $\boldsymbol{\beta_{lunemp*}(\tau)}$ | -0.45[c] (0.24) | -0.39[b] (0.19) | -0.31[c] (0.16) | -0.40[b] (0.17) | -0.41[b] (0.16) | -0.45[a] (0.16) | -0.22[c] (0.12) | -0.20[c] (0.11) | -0.33[a] (0.11) |
| $\boldsymbol{\beta_{lexc*}(\tau)}$ | 0.58[a] (0.15) | 0.47[a] (0.11) | 0.44[a] (0.10) | 0.56[a] (0.11) | 0.62[a] (0.09) | 0.56[a] (0.10) | 0.45[a] (0.07) | 0.50[a] (0.07) | 0.70[a] (0.08) |
| $\boldsymbol{\omega_0(\tau)}$ | 0.17[a] (0.04) | 0.15[a] (0.03) | 0.15[a] (0.03) | 0.16[a] (0.02) | 0.19[a] (0.02) | 0.17[a] (0.02) | 0.22[a] (0.03) | 0.27[a] (0.03) | 0.26[a] (0.04) |
| $\boldsymbol{\omega_1(\tau)})$ | 0.05 (0.05) | 0.01 (0.03) | 0.03 (0.03) | 0.05[c] (0.03) | 0.08[a] (0.02) | 0.08[a] (0.02) | 0.07[b] (0.03) | 0.07[b] (0.03) | 0.09[b] (0.04) |
| $\boldsymbol{\omega_2(\tau)}$ | 0.03 (0.04) | 0.06 (0.03) | 0.09[a] (0.03) | 0.07[a] (0.02) | 0.07[a] (0.02) | 0.07[a] (0.02) | 0.05[c] (0.03) | 0.04 (0.03) | 0.00 (0.04) |

*Notes*: Standard errors are reported in parentheses. Superscripts a, b, and c denote statistical significance at the 1%, 5%, and 10% levels, respectively. Lag length selection is determined according to the Akaike Information Criterion (AIC). The optimal lag structure of the model has been determined as (1,0,0,0,0,3) based on the AIC criterion (evaluated over 12,500 models).

*Source*: Author's own calculation.

As stated above, in Model 1 the dependent variable is linf**,** while the independent variables are lwsi**,** loil, lgdp, and lunemp. Initially, the calculated F-statistic is 28.64 which is greater than the upper critical values (I(1)) shown in Table 3. This result indicates the presence of cointegration among the variables. The long run results indicate that wages exert a consistently positive and statistically significant effect on inflation across all quantiles, confirming the presence of a wage–price spiral. At lower quantiles (τ=0.10–0.30), the coefficients are relatively high, suggesting that wage increases have a strong inflationary impact even in low-inflation regimes. In the middle quantiles (τ=0.40–0.60), the effect remains positive and significant, though slightly moderated, implying a steady pass-through of wage pressures into consumer prices under average inflation conditions. At the upper quantiles (τ=0.70–0.90), the positive impact persists but the magnitude decreases somewhat, which may reflect saturation effects or the influence of policy interventions aimed at curbing inflation when it reaches high levels. Overall, the results confirm that rising wages consistently contribute to higher inflation in all quantiles, but the strength of this effect changes depending on whether inflation is low, moderate, or high. The findings align with the evidence reported by Başkaya and Özmen (2013) as well as Kocaman and Biçerli (2019).

Moreover, the results demonstrate that the influence of oil prices on inflation is heterogeneous across the conditional distribution of inflation. At lower quantiles (τ=0.10–0.30), oil prices show no statistically significant effect, suggesting that during low-inflation periods, fluctuations in global energy markets exert only a limited pass-through to domestic prices. Beginning at the middle quantiles (τ=0.40–

0.60), the coefficients turn positive and statistically significant, indicating that oil price increases start to transmit more strongly into consumer prices as inflation rises toward its median levels. This impact becomes particularly pronounced at the upper quantiles (τ=0.70–0.90), where both the magnitude and significance of the coefficients increase sharply, underscoring the role of oil price shocks as a key inflationary driver in high-inflation regimes.

Furthermore, the results provide strong evidence for a growth-led disinflation effect. Across all significant quantiles, a 1% increase in GDP is associated with a decrease in inflation, though the magnitude of this effect varies across the distribution. At the lower quantiles (τ=0.10–0.20), the coefficients are small and statistically insignificant, suggesting that GDP growth does not meaningfully affect inflation in low-inflation regimes. From the middle quantiles (τ=0.30–0.60), the coefficients become negative and significant (around –0.5), indicating that a 1% rise in GDP reduces inflation by about 0.5% under moderate inflation conditions. The effect intensifies further in the upper quantiles (τ=0.70–0.90), where a 1% increase in GDP reduces inflation by as much as 0.8%.

In terms of the unemployment, the results provide evidence of a nonlinear Phillips curve relationship between unemployment and inflation. Across all quantiles, the coefficients are negative, indicating that higher unemployment tends to reduce inflation. This effect is strongest and most statistically significant in the lower and middle quantiles (τ=0.10–0.60), where a 1% increase in unemployment is associated with a relatively large decline in inflation, suggesting that the labour market slack plays an important disinflationary role in low-to-moderate inflation regimes. At the upper quantiles (τ=0.70–0.90), the coefficients remain negative, but their magnitude decreases and their significance weakens, implying that in high-inflation environments, the dampening effect of unemployment on inflation becomes less pronounced. This pattern suggests that while unemployment generally exerts downward pressure on prices, its role is more influential when inflation is low or moderate, whereas in high-inflation regimes other forces – such as wage–price spirals, exchange rate pass-through, or supply-side shocks – may dominate inflation dynamics.

Finally, the results reveal a strong and persistent positive relationship between exchange rates and inflation in all quantiles, confirming the presence of significant exchange rate pass-through effects in Türkiye. At the lower quantiles (τ=0.10–0.30), a 1% depreciation in the exchange rate increases inflation by roughly 0.44–0.58%, while at the middle quantiles (τ=0.40–0.60) this effect strengthens to about 0.56–0.63%, indicating that the inflationary impact of currency depreciation becomes more pronounced as inflation rises toward its median levels. At the upper quantiles (τ=0.70–0.90), the effect remains positive and significant, reaching as high as 0.71 at τ=0.90, which suggests that in high-inflation regimes exchange rate shocks act as a major driver of price increases. Overall, these findings underscore the economy's vulnerability to exchange rate fluctuations and highlight the importance of ensuring exchange rate stability and effective monetary–fiscal policy coordination to mitigate the inflationary consequences of currency depreciation.

The short-run results indicate that the contemporaneous effect of the exchange rate ($\omega_0$) is positive and statistically significant across all quantiles, confirming that exchange rate depreciations immediately translate into higher inflation. The first lag ($\omega_1$) is also generally positive and becomes statistically significant at the middle and upper quantiles (τ=0.50–0.90), suggesting that the impact of exchange rate shocks on inflation persists beyond the current period and continues into the following quarter. On the contrary, the second lag ($\omega_2$) is relatively small and mostly insignificant, implying that the pass-through effect weakens substantially over time and is largely absorbed within the first two periods. Overall, these results highlight that short-run exchange rate pass-through is both immediate and persistent in the near term, but its influence diminishes quickly thereafter.

The error correction term $\rho(\tau)$ coefficients are negative and statistically significant across all quantiles, confirming the existence of a long-run cointegration relationship and indicating that short-run deviations from equilibrium are systematically corrected over time. At the lower quantiles (τ=0.10–0.30), the adjustment speed ranges between 8% and 11%, suggesting a relatively slower correction process in low-inflation regimes. At the middle quantiles (τ=0.40–0.60), the coefficients increase slightly to around 11–12%, implying a moderate pace of adjustment. At the upper quantiles (τ=0.70–0.90), the coefficients reach their highest values, between 15% and 17%, which indicates that in high-inflation regimes, the system corrects disequilibria more rapidly. Overall, the results highlight that the speed of adjustment is not constant across the distribution, but rather accelerates under high-inflation conditions, reflecting stronger corrective dynamics when price pressures are elevated.

After obtaining the coefficients, the presence of asymmetry was tested using the Wald test, and the results are presented in Table 3.

**Table 3. Wald test results for model 1**

| Variables | $X^2$ stat. | p-value |
|---|---|---|
| $c_0$ | 3.100 | 0.684 |
| $\rho(\tau)$ | 3.402 | 0.492 |
| $\beta_{lwsi*}(\tau)$ | 8.306 | 0.080 |
| $\beta_{loil*}(\tau)$ | 1.573 | 0.813 |
| $\beta_{lgdp*}(\tau)$ | 4.010 | 0.404 |
| $\beta_{lunemp*}(\tau)$ | 3.196 | 0.525 |
| $\beta_{lexc*}(\tau)$ | 3.356 | 0.500 |
| $\omega_0(\tau)$ | 3.487 | 0.479 |
| $\omega_1(\tau)$ | 2.012 | 0.733 |
| $\omega_2(\tau)$ | 5.149 | 0.272 |

*Source*: Author's own calculation.

The Wald test results indicate that, for all variables except wages, the null hypothesis of coefficient equality across quantiles cannot be rejected, suggesting that their effects on inflation are broadly homogeneous across different inflation regimes. This implies that shocks to exchange rates, oil, GDP, and unemployment exert relatively stable influences on inflation regardless of whether the economy is in a low-, medium-, or high-inflation regimes. By contrast, the wage variable exhibits significant heterogeneity across quantiles, meaning that its impact on inflation varies with the level of inflation. In particular, wage shocks exert disproportionately stronger effects in higher inflation regimes, which is consistent with the wage–price spiral literature and highlights the direction-dependent role of wages in shaping price dynamics. From a policy perspective, this finding underscores the importance of carefully monitoring wage developments in inflation-targeting frameworks. While stabilising other macroeconomic variables may help anchor prices uniformly across regimes, wage dynamics appear to be more volatile and context-dependent, requiring stronger coordination between monetary policy, income policies, and labour market institutions to prevent wage shocks from fuelling persistent inflationary pressures.

Having discussed the empirical results of the first model, the analysis now proceeds to the second model, where the QARDL framework is employed to examine the impact of inflation on wages across different quantiles. Therefore, in Model 2 the dependent variable is lwsi**,** while the independent variables are linf**,** loil, lgdp, and lunemp. Table 4 provides the results of the cointegration bounds test and subsequently reports the QARDL model estimations for Model 2.

**Table 4. Quantile ARDL results for model 2**

| **Panel A: Bounds test statistics** | | | | | | |
|---|---|---|---|---|---|---|
| 1% | | 5% | | 10% | | F stat: 8.424 |
| I(0) | I(1) | I(0) | (I1) | (I0) | I(1) | |
| 3.06 | 4.15 | 2.39 | 3.38 | 2.08 | 3 | |

| **Panel B: Quantile ARDL** | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | | | ***Quantiles*** | | | | | |
| Variables | 0.10 | 0.20 | 0.30 | 0.40 | 0.50 | 0.60 | 0.70 | 0.80 | 0.90 |
| $c_1(\tau)$ | -12.5$^{a}$ (0.93) | -12.2$^{a}$ (0.96) | -12.3$^{a}$ (1.22) | -13.5$^{a}$ (1.68) | -14.5$^{a}$ (2.12) | -15.5$^{a}$ (2.26) | -15.8$^{a}$ (2.28) | -17.8$^{a}$ (2.71) | -17.4$^{a}$ (1.92) |
| $\rho(\tau)$ | -0.14$^{a}$ (0.03) | -0.13$^{a}$ (0.02) | -0.10$^{a}$ (0.02) | -0.08$^{a}$ (0.02) | -0.07$^{a}$ (0.01) | -0.07$^{a}$ (0.01) | -0.07$^{a}$ (0.01) | -0.07$^{a}$ (0.01) | -0.09$^{a}$ (0.01) |
| $\beta_{linf*}(\tau)$ | 0.87$^{a}$ (0.11) | 0.71$^{a}$ (0.13) | 0.78$^{a}$ (0.16) | 0.72$^{a}$ (0.21) | 0.74$^{a}$ (0.25) | 0.85$^{a}$ (0.23) | 0.94$^{a}$ (0.22) | 0.86$^{a}$ (0.23) | 1.17$^{a}$ (0.16) |
| $\beta_{lexc*}(\tau)$ | 0.07 (0.62) | 0.20 (0.12) | 0.12 (0.15) | 0.13 (0.20) | 0.05 (0.23) | -0.04 (0.22) | -0.12 (0.22) | -0.10 (0.22) | -0.39 (0.17) |
| $\beta_{lgdp*}(\tau)$ | 1.30$^{a}$ (0.10) | 1.25$^{a}$ (0.10) | 1.26$^{a}$ (0.13) | 1.38$^{a}$ (0.18) | 1.51$^{a}$ (0.24) | 1.57$^{a}$ (0.24) | 1.53$^{a}$ (0.23) | 1.70$^{a}$ (0.27) | 2.21$^{a}$ (0.47) |
| $\beta_{loil*}(\tau)$ | -0.15$^{a}$ (0.05) | -0.07 (0.05) | -0.12 (0.07) | -0.11 (0.09) | -0.19$^{c}$ (0.11) | -0.26$^{b}$ (0.11) | -0.29$^{b}$ (0.11) | -0.24$^{b}$ (0.11) | -0.29$^{a}$ (0.08) |
| $\beta_{lunemp*}(\tau)$ | -0.54$^{a}$ (0.20) | -0.27 (0.20) | -0.30 (0.25) | -0.35 (0.33) | -0.59 (0.42) | -0.51 (0.41) | -0.26 (0.37) | -0.32 (0.38) | 0.13 (0.26) |
| $\varphi_1(\tau)$ | -0.05 (0.09) | 0.07 (0.07) | 0.08 (0.07) | 0.02 (0.07) | 0.04 (0.07) | 0.06 (0.07) | 0.04 (0.07) | 0.03 (0.08) | -0.06 (0.09) |
| $\varphi_2(\tau)$ | 0.24$^{a}$ (0.08) | 0.36$^{a}$ (0.07) | 0.35$^{a}$ (0.06) | 0.34$^{a}$ (0.06) | 0.27$^{a}$ (0.06) | 0.26$^{a}$ (0.06) | 0.26$^{a}$ (0.07) | 0.22$^{a}$ (0.07) | 0.23$^{a}$ (0.08) |
| $\varphi_3(\tau)$ | 0.05 (0.08) | 0.15$^{b}$ (0.07) | 0.19$^{a}$ (0.06) | 0.14$^{b}$ (0.06) | 0.10$^{c}$ (0.06) | 0.10$^{c}$ (0.06) | 0.12$^{b}$ (0.06) | 0.18$^{a}$ (0.06) | 0.29$^{a}$ (0.08) |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| $\sigma_0(\tau)$ | 0.30[b] (0.12) | 0.13 (0.10) | 0.14 (0.09) | 0.12 (0.09) | 0.11 (0.09) | 0.10 (0.09) | 0.17[c] (0.09) | 0.13 (0.10) | 0.30[b] (0.11) |
| $\sigma_1(\tau)$ | 0.23[c] (0.12) | 0.13 (0.10) | 0.06 (0.09) | 0.13 (0.09) | 0.11 (0.09) | 0.15 (0.09) | 0.16[c] (0.09) | 0.19[c] (0.10) | 0.32[a] (0.12) |
| $\zeta_0(\tau)$ | -0.31[b] (0.14) | -0.28[b] (0.12) | -0.23[b] (0.11) | -0.20[c] (0.10) | -0.21[c] (0.10) | -0.15 (0.11) | -0.09 (0.11) | -0.03 (0.12) | 0.02 (0.14) |
| $\zeta_1(\tau)$ | 0.13 (0.15) | 0.20[c] (0.12) | 0.16 (0.11) | 0.17 (0.11) | 0.22[b] (0.11) | 0.19[c] (0.11) | 0.10 (0.11) | 0.08 (0.12) | 0.01 (0.14) |

*Notes*: Standard errors are reported in parentheses. Superscripts a, b, and c denote statistical significance at the 1%, 5%, and 10% levels, respectively. Lag length selection is determined according to the Akaike Information Criterion (AIC). The optimal lag structure of the model has been determined as (4, 2, 0, 0, 0, 2) based on the AIC criterion (evaluated over 12,500 models).

*Source*: Author's own calculation.

The computed F-statistic, equal to 8.424, surpasses the upper critical value (I(1)) reported in Table 3, providing robust evidence of a cointegrating relationship among the variables. The long run results indicate that inflation (linf) has a consistently positive and significant effect on wages across all quantiles, though the strength of this relationship varies along the wage distribution. At the lower quantiles (τ=0.10–0.30), a 1% increase in inflation raises wages by about 0.72 to 0.88%, suggesting a strong but slightly less-than-proportional adjustment when wages are relatively low. In the middle quantiles (τ=0.40–0.60), the coefficients remain positive and stable (around 0.72–0.85), reflecting a steady inflation–wage pass-through under average wage conditions. However, in the upper quantiles (τ=0.70–0.90), however, the coefficients become much larger, reaching their maximum at 1.17 at τ=0.90, which indicates that wage growth outpaces inflation in the higher segments of the distribution. This heterogeneous pattern highlights the presence of a wage–price spiral dynamic: as inflation rises, wages respond positively, and in higher wage values, they may even overshoot inflation, potentially feeding further into price pressures. The results suggest that while inflation pass-through to wages is evident across the entire distribution, it is particularly strong in higher wage values, consistent with stronger bargaining power and indexation practices. From a policy perspective, this underscores the risk that persistent inflation can fuel a reinforcing cycle of rising wages and prices, complicating stabilisation efforts. In the first model, we observed that wages exert an influence on inflation, whereas in this model, inflation is found to affect wages. These findings support the existence of a wage–price spiral and are consistent with the results reported by Hoxha (2010) and Özer and Gülşen (2025).

In terms of the exchange rate, in the lower quantiles (τ=0.10–0.40), the coefficients are positive but statistically insignificant, indicating that exchange rate fluctuations do not exert a meaningful effect on wages when wage levels are relatively low to moderate. At the middle quantiles (τ=0.50–0.60), the coefficients turn slightly negative but remain insignificant, suggesting weak and unstable effects. At the higher quantiles (τ=0.70–0.90), the coefficients become increasingly negative (reaching –0.39 at τ=0.90), though still statistically insignificant given the relatively large standard errors. In general, these results suggest that exchange rate movements do not have a systematic or robust impact on wages across the distribution. While there is a tendency for exchange rate depreciations to exert downward pressure on wages in the higher quantiles, the lack of statistical significance implies that this relationship is weak and not persistent.

The estimated coefficients of lgdp (real GDP) on wages at the lower quantiles (τ=0.10–0.30), the coefficients are around 1.25–1.30, implying that a 1% increase in GDP is associated with more than a proportional increase in wages among lower-wage values. In the middle quantiles (τ=0.40–0.60), the coefficients rise further (1.38–1.57), indicating a stronger wage response to GDP in average wage levels. At the higher quantiles (τ=0.70–0.90), the coefficients become even higher, reaching 2.21 at τ=0.90. This suggests that in the upper segments of the wage distribution, wage growth is highly responsive to output growth, with GDP expansions generating more than twice the percentage increase in wages. Overall, the results show that economic growth exerts a strong and heterogeneous positive effect on wages, with the impact being especially pronounced at the higher quantiles.

In terms of the oil prices, the results indicate that oil price shocks exert a consistently negative impact on wages across the distribution, with the effect becoming stronger and more statistically robust in the middle and upper quantiles. While the influence is modest and less consistent in the lower wage values, it intensifies in the median and higher wage segments, where a 1% increase in oil prices reduces wages by up to 0.29%. This pattern suggests that oil price increases operate through cost-push channels and raising production costs. Although it is expected that rising energy prices (a negative supply shock) will increase inflation through the cost-push channel, leading workers to demand higher wages to compensate for the loss in their real wages, the findings indicate that workers' bargaining power is not strong in the face of such shocks, especially in higher inflation regimes.

Finally, in terms of the unemployment, the results indicate that unemployment has a statistically significant negative impact on wages only at the lower quantile (τ=0.10), where a 1% increase in unemployment reduces wages by about 0.54%. At higher quantiles, although the coefficients remain mostly negative, they are statistically insignificant, implying that unemployment does not exert a robust effect on middle- and high-wage values.

The short-run results indicate that immediate wage effects are mostly insignificant, while the second lag shows a strong and significant positive impact across nearly all quantiles. The third lag becomes significant only at higher quantiles, suggesting some persistence in wage dynamics, particularly among upper wage values. Moreover, current inflation $\sigma_0(\tau)$ has a positive and significant effect on wages at the lower (τ=0.10) and upper (τ=0.70, 0.90) quantiles, suggesting stronger pass-through in low- and high-wage values. The lagged inflation term $\sigma_1(\tau)$is also positive and significant mainly at higher quantiles, indicating that wage adjustments to inflation persist over time and are more pronounced among upper wage values. Overall, inflation shocks raise wages in the short run, with effects strongest in the tails of the wage distribution. Furthermore, current unemployment changes $\zeta_0(\tau)$

have a negative and significant effect on wages at lower quantiles (τ=0.10–0.30), indicating that rising unemployment quickly reduces wages among low-wage values. The lagged unemployment term $\zeta_1(\tau)$turns positive and significant at the middle quantiles (τ=0.40–0.60), suggesting some delayed upward adjustment of wages once unemployment pressures ease. Overall, the impact of unemployment on wages is strongest and negative in the lower quantiles, while weaker and less consistent in the middle and upper parts of the distribution. The error correction term $\rho(\tau)$ coefficients are negative and statistically significant, confirming the existence of a stable long-run relationship and the adjustment of short-run deviations back toward equilibrium. The speed of adjustment varies across quantiles, ranging from about –0.07 to –0.15. The adjustment is relatively faster at the lower and upper quantiles (τ=0.10 and τ=0.90), suggesting that wages return to long-run equilibrium more quickly when they are at the extremes of the distribution. Overall, the results highlight a stable and significant error-correcting mechanism. In the next step, the Wald test is conducted. Table 5 presents Model 2's Wald test results.

**Table 5. Wald test results for model 2**

| Variables | $X^2$ stat. | p-value |
|---|---|---|
| $c_1(\tau)$ | 3.985 | 0.40 |
| $\rho(\tau)$ | 2.251 | 0.68 |
| $\beta_{linf*}(\tau)$ | 2.582 | 0.62 |
| $\beta_{loil*}(\tau)$ | 3.002 | 0.08 |
| $\beta_{lgdp*}(\tau)$ | 15.39 | 0.00 |
| $\beta_{lunemp*}(\tau)$ | 0.732 | 0.94 |
| $\beta_{lexc*}(\tau)$ | 1.747 | 0.78 |
| $\varphi_1(\tau)$ | 0.480 | 0.97 |
| $\varphi_2(\tau)$ | 0.891 | 0.92 |
| $\varphi_3(\tau)$ | 1.075 | 0.89 |
| $\sigma_0(\tau)$ | 3.184 | 0.52 |
| $\sigma_1(\tau)$ | 3.872 | 0.42 |
| $\zeta_0(\tau)$ | 1.175 | 0.88 |
| $\zeta_1(\tau)$ | 4.809 | 0.30 |

*Source*: Author's own calculation.

As indicated above, in the Wald test, the null hypothesis imposes coefficient equality across quantiles ($H_0$: $\beta_{0.10} = \beta_{0.20} = \ldots = \beta_{0.90}$). Although the estimated coefficients in the lwsi model differ in magnitude across quantiles, the Wald test does not reject the null for most of the variables.

While the QARDL framework provides reliable estimates of the short- and long-run dynamics of the price–wage spiral in Türkiye, it does not allow for direct inference of causal relationships. Therefore, to rigorously examine the causal linkages among linf, lwsi, loil, lunemp, lexc and lgdp, the Fourier Toda–Yamamoto (F-TY) causality test is employed, and the corresponding results are presented in Table 6.

**Table 6. Fourier causality test results**

| Causality | Wald-stat. | Asymp. p-value | Bootstrap p-value | k | p |
|---|---|---|---|---|---|
| lwsi→linf | 5.570[b] | 0.018 | 0.026 | 1 | 1 |
| lexc→linf | 30.536[a] | 0.000 | 0.000 | 1 | 5 |
| linf→lexc | 9.469 | 0.092 | 0.106 | 2 | 5 |
| lgdp→linf | 0.359 | 0.549 | 0.535 | 1 | 1 |
| linf→lgdp | 0.182 | 0.670 | 0.666 | 1 | 1 |
| lunemp→linf | 0.963 | 0.618 | 0.613 | 1 | 2 |
| linf→lunemp | 1.278 | 0.528 | 0.526 | 1 | 2 |
| loil→linf | 1.422 | 0.491 | 0.479 | 1 | 2 |
| linf→loil | 1.765 | 0.414 | 0.414 | 1 | 2 |
| linf→lwsi | 19.915[a] | 0.000 | 0.000 | 1 | 1 |
| lexc→lwsi | 10.947 | 0.090 | 0.117 | 1 | 6 |
| lwsi→lexc | 8.702 | 0.191 | 0.210 | 1 | 6 |
| lgdp→lwsi | 12.196[c] | 0.058 | 0.085 | 1 | 6 |
| lwsi→lgdp | 7.592 | 0.270 | 0.272 | 1 | 6 |
| lunemp→lwsi | 1.022 | 0.600 | 0.584 | 1 | 2 |
| lwsi→lunemp | 1.777 | 0.411 | 0.410 | 1 | 2 |
| loil→lwsi | 2.517 | 0.642 | 0.610 | 1 | 4 |
| lwsi→loil | 1.465 | 0.833 | 0.813 | 1 | 4 |

*Notes*: Superscripts a, b, and c denote statistical significance at the 1%, 5%, and 10% levels, respectively. The maximum order of integration (dmax) is identified as 1 for all variables. The maximum lag length (p) is set to 6, with the Akaike Information Criterion employed to select the optimal lag structure. The maximum frequency (k) is specified as 3, and its optimal value is determined through the estimation procedure. The analysis is conducted using GAUSS codes developed by Nazlıoğlu (2021), and bootstrap p-values are computed based on 10,000 replications.

*Source*: Author’s own calculation.

The F-TY causality test results reveal a strong bidirectional relationship between wages and inflation, as both directions (linf → lwsi and lwsi → linf) are statistically significant, providing clear evidence of a wage–price spiral. In addition, a unidirectional causality is found from exchange rates to inflation, confirming the presence of exchange rate pass-through effects. Economic growth also shows a weak but significant causal influence on wages at the 10% level, consistent with productivity–wage linkages. By contrast, no significant causal relationships are observed for unemployment or oil prices, suggesting that their effects on wages and inflation are likely indirect. Overall, these findings highlight that inflationary dynamics in Türkiye are primarily shaped by the interplay of wages and exchange rates, while growth plays a secondary role and other macroeconomic variables exert more limited direct influence.

## 5. Concluding remarks

This study investigates whether the wage–price spiral holds in the case of the Turkish economy. Employing the Quantile Autoregressive Distributed Lag (QARDL) approach, the analysis explores the impact of wage increases on inflation across different inflation regimes, while also assessing how inflation affects wages at different quantiles. To strengthen the empirical framework, energy prices, the exchange rate, unemployment, and real GDP are incorporated as control variables.

The obtained results of the first model show that wage increases exert a consistent upward effect on inflation across all quantiles, although the magnitude of this relationship varies depending on whether the economy is in a low, moderate, or high inflation regime. The Wald test further confirms the asymmetric impact of wage increases on inflation. Moreover, the findings indicate that the impact of oil prices on inflation is heterogeneous across the inflation distribution: while insignificant during low-inflation periods, the effect becomes significant from the median levels onward and is particularly pronounced in high-inflation regimes. From a policy perspective, these findings highlight the asymmetric vulnerability of the Turkish economy to oil price fluctuations. While low-inflation environments appear relatively resilient, periods of elevated inflation amplify the pass-through of energy price shocks, complicating the stabilisation role of monetary policy. the results suggest that central banks in energy-importing economies face greater challenges in curbing inflationary pressures when inflation is relatively high. Thus, coordinated policy responses – combining prudent monetary policy with structural measures such as enhancing energy efficiency, diversifying the energy mix, and strengthening institutional frameworks – are crucial to mitigate the inflationary impact of oil price shocks and preserve macroeconomic stability especially in high inflation environment. Furthermore, the findings offer robust evidence to support a growth-driven disinflation effect. Across all quantiles, a 1% rise in GDP is linked to a reduction in inflation, although the strength of this effect differs across the distribution. Moreover, the results provide evidence of a nonlinear Phillips curve relationship between unemployment and inflation. Although unemployment typically places downward pressure on prices, its influence is stronger in low- and moderate-inflation regimes, while in high-inflation regimes other forces tend to dominate inflation dynamics. Finally, the results indicate a strong and persistent positive link between exchange rates and inflation across all quantiles, confirming substantial exchange rate pass-through in Türkiye. In high-inflation regimes, exchange rate shocks emerge as a key driver of price increases. Taken together, these findings emphasise the economy's susceptibility to exchange rate fluctuations and highlight the critical importance of exchange rate stability and effective coordination between monetary and fiscal policies in mitigating the inflationary effects of currency depreciation.

Second model's results show that although inflation pass-through to wages is observed across the entire distribution, it is especially pronounced in the upper wage quantiles, reflecting stronger bargaining power and indexation mechanisms. From a policy standpoint, this highlights the risk that persistent inflation may trigger a self-reinforcing cycle of wage and price increases, thereby complicating stabilisation efforts. Moreover, exchange rate fluctuations do not exert a statistically significant influence on wages across the distribution. Furthermore, the results indicate that increases in real GDP (economic growth) exert a strong yet heterogeneous positive influence on wages, with the effect particularly evident in the upper quantiles. With respect to oil prices, the results show that oil price shocks exert a consistently negative influence on wages across the distribution, with the effect becoming stronger and more statistically robust at the middle and upper quantiles. Although rising energy prices—as a negative supply shock—are typically expected to fuel inflation through the cost-push channel and prompt workers to demand higher wages to offset real income losses, the findings suggest that workers' bargaining power remains weak in the face of such shocks, particularly under high-inflation regimes. From a policy standpoint, these results underscore the vulnerability of labour incomes to external energy shocks in an energy-importing economy such as Türkiye. They further highlight the need for policies that reduce dependence on imported oil, improve energy efficiency, and provide targeted income-support measures to mitigate the adverse distributional effects of oil price volatility on workers' welfare. Finally, the results show that unemployment significantly reduces wages only at the lower quantile ($\tau$=0.10), while its impact in the middle and upper quantiles is negative but statistically insignificant.

The Fourier Toda–Yamamoto (F-TY) causality test further reinforces these findings by revealing a statistically significant bidirectional causality between wages and inflation, thereby providing robust evidence of a wage–price spiral in the Turkish economy.

In light of these findings, it is evident that breaking the wage–price spiral in Türkiye requires more than temporary measures. Persistent inflation, if left unchecked, risks reinforcing cycles of wage and price increases that complicate stabilisation efforts. Policymakers should therefore prioritise credible disinflation strategies, strengthen expectation management, and design wage-setting frameworks that balance the protection of workers' real incomes with the need for macroeconomic stability. At the same time, reducing external vulnerabilities – particularly dependence on imported energy and exposure to exchange rate shocks – remains crucial. Without such a comprehensive and coordinated policy approach, the wage–price spiral may continue to pose a serious threat to long-term price stability and sustainable growth.

## References

[1] Abdioğlu, Z. (2015), *Ücret-Fiyat Spirali Dinamiklerinin Asimetrik Etkisi: Türkiye Örneği. Marmara Üniversitesi İktisadi Ve İdari Bilimler Dergisi*, 36(1), 245-268, https://doi.org/10.14780/iibdergi.201417547.

[2] Akgül, O., Bükey, A.M. (2020), *Türkiye'de enflasyon ile asgari ücretler arasındaki ilişki ve ücret-fiyat sarmalı*. *Journal of Social Policy Conferences*, 78, 257-282, Istanbul University, Turkiye, https://doi.org/10.26650/jspc.2020.78.0026.

[3] Alvarez, J., Bluedorn, J.C., Hansen, N.J., Huang, Y., Pugacheva, E., Sollaci, A. (2024), *Wage–price spirals: what is the historical evidence?*. *Economica*, 91(364), 1291-1319, https://doi.org/10.1111/ecca.12543.

[4] Başkaya, Y.S., Özmen, M.U. (2013), *Türkiye'de Asgari Ücret-ÜFE Enflasyonu İlişkisi Üzerine Ampirik Bir analiz, Ekonomi Notları,* (23), Türkiye Cumhuriyet Merkez Bankası, 6 Eylül, Ankara, Turkiye, https://www.tcmb.gov.tr.

[5] Bernanke, B., Blanchard, O. (2025), *What caused the US pandemic-era inflation?*. *American Economic Journal: Macroeconomics*, 17(3), 1-35, https://doi.org/10.1257/mac.20230195.

[6] Bicerli, M.K., Kocaman, M. (2019), *The impact of minimum wage on unemployment, prices, and growth: A multivariate analysis for Turkey*. *Economic Annals*, 64(221), 65-84, https://doi.org/10.2298/eka1921065k.

[7] Blanchard, O.J. (1986), *The wage price spiral*. *The Quarterly Journal of Economics*, 101(3), 543-565, https://doi.org/10.2307/1885696.

[8] Cao, H., Khan, F.U., Shaikh, P.A., Khan, F., Zehri, M., Yousaf, H. (2025), *Does oil price shock drive inflation? Evidence from G20 countries*. *Journal of the Knowledge Economy*, 16(1), 904-924, https://doi.org/10.1007/s13132-024-01877-1.

[9] Cho, J.S., Kim, T.-H., Shin, Y. (2015), *Quantile cointegration in the autoregressive distributed-lag modeling framework*. *Journal of Econometrics,* 188(1), 281-300, https://doi.org/10.1016/j.jeconom.2015.05.003.

[10] Dickey, D.A., Fuller, W.A. (1981), *Likelihood ratio statistics for autoregressive time series with a unit root*. *Econometrica: journal of the Econometric Society*, 1057-1072, https://doi.org/10.2307/1912517.

[11] Hoxha, A. (2010), *Causality between Prices and Wages: VECM Analysis for EU-27*. *Romanian Economic Journal*, 13(37).

[12] Lütkepohl, H. (2005), *New Introduction to Multiple Time Series Analysis*. Springer-Verlag, Berlin Heidelberg, Germany, https://doi.org/10.1007/978-3-540-27752-1.

[13] Nazlioglu, S. (2021). TSPDLIB: GAUSS Time Series and Panel Data Methods (Version 2.0)—Source code. https://github.com/aptech/tspdlibNazlioglu.

[14] Nazlioglu, S., Gormus, N.A., Soytas, U. (2016), *Oil prices and real estate investment trusts (REITs): Gradual-shift causality and volatility transmission analysis*. *Energy economics*, 60, 168-175, https://doi.org/10.1016/j.eneco.2016.09.009.

[15] Özer, M.O., İrteş Gülşen, M. (2025), *Türkiye'de Asgari Ücret ile Enflasyon, İşsizlik ve Büyüme Oranları Arasındaki İlişki*. *Erzurum Teknik Üniversitesi Sosyal Bilimler Enstitüsü Dergisi,* (21), 71-84, https://doi.org/10.29157/etusbed.1501796.

[16] Özyiğit, M., Çelik, A., Petek, A. (2019), *Türkiye'de 1980 Sonrası Reel Asgari Ücretin Belirleyicileri*. *Çalışma Ve Toplum*, 2(61), 1029-1068.

[17] Perron, P. (1989), *The great crash, the oil price shock, and the unit root hypothesis. Econometrica: Journal of the Econometric Society*, 1361-1401, https://doi.org/10.2307/1913712.

[18] Saunders, M., Denniss, R. (2022), *Wage price spiral or price wage spiral. The Australia Institue-Research that matters*. (Electronic paper), https://australiainstitute.org.au/wp-content/uploads/2022/05/220517-Price-Wage-Spiral.pdf.

[19] Toda, H.Y., Yamamoto, T. (1995), *Statistical inference in vector autoregressions with possibly integrated processes. Journal of econometrics*, 66(1-2), 225-250, https://doi.org/10.1016/0304-4076(94)01616-8.

[20] Zeira, J. (1989), *Inflationary inertia in a wage-price spiral model. European Economic Review*, 33(8), 1665-1683, https://doi.org/10.1016/0014-2921(89)90085-8.