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# Testing the Long-Run Relationship Between Oil Price and Inflation in Tunisia: A Fourier CUSUM Cointegration Test

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This study examines the relationship between oil prices and inflation in Tunisia using data from 1985 to 2024. Applying the innovative Fourier CUSUM test, the findings indicate no long-run connection between the two variables, suggesting that subsidy policies effectively buffer Tunisia from global oil price shocks. However, QARDL results reveal short-term asymmetries: oil price increases cause temporary rises in inflation, while natural gas prices have a dampening effect. Overall, these results highlight the stabilizing - yet fiscally risky - role of subsidies, offering valuable insights for energy-importing economies.

## I. Introduction

Numerous studies have demonstrated a long-term relationship between oil prices and inflation in various economies. Research conducted in India (Sultan et al., 2020), South Africa (Niyimbanira, 2013), and Asia-Pacific countries (Jiranyakul, 2021) has confirmed this connection using cointegration techniques. Generally, oil price shocks tend to increase inflation, although the extent and duration of this effect can differ. These studies reveal that oil price fluctuations influence inflation both in the short and long term, with the impact being especially marked during periods of high and volatile oil prices. However, the pass-through from oil prices to consumer prices is typically partial, varying across countries and time periods (Renou-Maissant, 2019). Notably, this relationship can be moderated by active fiscal stabilization policies. In line with price insulation theories (Gelos & Ustyugova, 2017), governments may deploy subsidy regimes to shield domestic economies from global energy shocks, as evidenced by Tunisia's hydrocarbon subsidies, which help disconnect local prices from international markets.

Structural breaks in oil price trends can also significantly affect this relationship. Major events such as the Asian financial crisis in 1997, the global financial crisis in 2008, the COVID-19 pandemic, and the Russia-Ukraine war have all influenced oil and gas prices, resulting in structural breaks within the data. Traditional cointegration studies with structural breaks often assume a known number of changes and employ dummy variables to capture unexpected shifts. In reality, time series data often contain several unknown structural breaks. Building on the work of Hao and Inder (1996), Xiao and Phillips (2002), and Tsong et al. (2016), this paper proposes a Fourier CUSUM cointegration test to better capture the essential features of structural changes and evaluate the long-run relationship between oil prices and inflation in Tunisia.

The findings of this study present a notable exception to the commonly observed oil-inflation nexus: cointegration tests reveal no long-run relationship between inflation and oil prices in Tunisia, a result that underscores the price insulation provided by subsidy policies. However, the QARDL estimation shows that increases in oil prices temporarily raise inflation in the short run, indicating incomplete pass-through effects. This duality illustrates the trade-offs of fiscal stabilization while subsidies protect households from price shocks, they also introduce long-term fiscal risks. The decoupling of long-term trends amid short-term volatility highlights the complex interplay between global energy markets and domestic policies in energy-importing nations.

The paper is organized as follows: Section II covers data and methods, Section III outlines results, and Section IV concludes.

## II. Data and Methodology

### A. Data

This study utilizes 476 monthly observations for Tunisia, covering the period from January 1985 to January 2024. Data are drawn from the U.S. Energy Information Administration, World Bank (WDI), and the Tunisian Institute of

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Statistics. The variables include the inflation rate (*inf*), WTI crude oil price (in USD per barrel) (*wti*), Henry Hub natural gas price (*gasp*), and the geopolitical uncertainty index (*gpr*), all expressed in natural logarithms and aligned with international standards.

## B. Methodology

(Hao & Inder, 1996) and (Xiao & Phillips, 2002) report that the null hypothesis of cointegration can be assessed by examining the variation in the residual process of cointegration regression with the CUSUM test for structural stability. The cointegration test described here builds on previous CUSUM tests by adding trigonometric Fourier terms to the cointegration regression as follows:

$$y_t = \alpha_0 + d(t) + \beta_t X_t + \mu_t,\ t = 1, \ldots, T \tag{1}$$

Where $d(t)$ is the deterministic term

$$d(t) = \theta_1 + \lambda_1 Sin(\frac{2\pi kt}{T}) + \lambda_2 Cos(\frac{2\pi kt}{T}) \tag{2}$$

Here, $k$ represents the frequency used to estimate the unknown number and locations of structural changes, $t$ is the trend term, and $T$ denotes the sample size. The optimal value of $k$ is determined using the AIC criterion within the interval $k = [0.01, \ldots, 3]$. Fractional frequencies are used to identify persistent (standing) breaks, while integer frequencies reveal temporary (transitory) breaks.

Residuals from Equation (1) are used to construct the following cumulated sum statistic:

$$CS_n = max_{k=1,\ldots,n} \frac{1}{\widehat{\sigma}\sqrt{T}} \left| \sum_{t=1}^{k} \widehat{\mu_t^+} \right| \tag{3}$$

where $\widehat{\sigma}$ is the estimated error variance and $\widehat{\mu}$ denotes residuals.

By following (Hao & Inder, 1996), we use a data-generation process from the following bivariate regression model:

$$y_t = \beta x_t + \varepsilon_t, \quad t = 1, \ldots, T \tag{4}$$
$$x_t = x_{t-1} + \nu_t$$

Here, $x_t$ is integrated of order one, $I(1)$, without drift. The critical values reported in Table 1 are obtained via direct simulation, using $n = 1,000$ with 10,000 replications for $k = 1,\ 2,\ 3$ frequencies and $p = 1,\ 2,\ 3,\ 4$ regressors. A Monte Carlo experiment was also conducted to examine the finite sample performance and power of the Fourier CUSUM cointegration test (see Table 2). In this experiment, we set $epsilon_t = gammaepsilon_{t-1} + zeta_t$. When $\gamma < 1$, $y_t$ and $x_t$ are cointegrated, whereas $\gamma = 1$ indicates no cointegration. The experiment considered the following sample sizes: $n = 100,\ 200,\ 300,\ 500$, with $\gamma = \ 0\ and\ 0.2$, each for 5,000 replications. The simulation results indicate that over rejection of the null hypothesis of cointegration is more likely for small $n$ and $k$, and increases as $\gamma$, $n$, and $k$ grow.

Table 3 presents simulation results on test power, indicating that power improves with increasing size and frequency.

The econometric paradigm involves three distinct stages. Initially, the stationarity of each variable is assessed using the Augmented Dickey-Fuller, Zivot-Andrews, and Fourier Augmented Dickey-Fuller unit root tests. If unit roots are present, the analysis proceeds with the Fourier CUSUM test and the cointegration tests as outlined by Tsong et al. (2016). In the final stage, the QARDL model is utilized for evaluation.

## III. Empirical Results

The initial step involves testing for stationarity using the standard ADF test, the Zivot-Andrews test (which accounts for structural breaks), and the Fourier ADF test (Enders & Lee, 2012). As presented in Panel of Table 4, the *F*-test identifies only *gasp* and *gpr* as significant at the 1% and 10% levels, respectively. Accordingly, the ADF test is applied to *gasp* and *gpr*, while the FADF test is used for *inf* and *wti*.

The FADF test does not support the presence of a unit root for *wti*, gasp, and *gpr*, while it does not reject the unit root for inf. All series are stationary in first differences. Likewise, the ADF test finds evidence against the null hypothesis for *gasp* and *gpr* at the 1% and 5% significance levels, respectively, but does not do so for *inf* and *wti*; all series show stationarity in their first differences. The Zivot-Andrews test indicates structural breaks in *gasp* and *gpr*, and finds no unit root for *inf* and *wti*.

Overall, the unit root tests reveal a mixed order of integration I(0) and I(1) with no evidence of I(2). This outcome makes cointegration testing necessary. As shown in Table 1 Panel B, the findings indicate no cointegration and, consequently, no long-run relationship between inflation, geopolitical uncertainty, and oil and gas prices in Tunisia.

Given the absence of a cointegrating relationship, the QARDL model proposed by Cho, Kim, and Shin (2015) is estimated. The QARDL model is specified as follows:

$$Y_t = \alpha(\tau) + \sum_{j=1}^{p} \phi_j(\tau) Y_{t-j} + \sum_{j=0}^{q} \gamma_j(\tau)' X_{t-j} + U_t(\tau) \tag{5}$$

where $Y_t$ represents the the consumer price index $\mathrm{inf}_t$, $X_t$ represents the oil price ($wti_t$), $gasp_t$ denotes gas price, and $gpr_{t}\ $represents geopolitical uncertainty index. $\tau = \{0.1, \ldots, 0.9\}$ donates the different quantiles. The optimal lags p and q are selected using AIC criterion. $\beta^* = \frac{\sum_{j=1}^{p} \phi_j(\tau)}{\sum_{j=1}^{q} \gamma_j(\tau)}$ is the long-run cointegration parameter. $\phi$ represents the total short-term impact of past variations in the dependent variable $\mathrm{Y}_t$ on its current variation. $\gamma$ gives the cumulative short-run effect of actual and past changes in the regressors on the current change in $\mathrm{Y}_{t.}$

The QARDL model results (Table 4, Panel C) indicate that there is no significant long-term relationship between inflation, geopolitical uncertainty, and oil and gas prices in Tunisia. This finding is consistent with outcomes observed in subsidy-dependent economies, where interventions such as fuel subsidies and price caps dampen the transmission of global oil shocks to domestic inflation.

In the short term, however, increases in oil prices trigger temporary spikes in inflation a typical pattern in energy-importing countries with incomplete pass-through. Sectors such as manufacturing, agribusiness, and logistics face rising input costs (Baumeister & Kilian, 2016), which are often passed on to consumers and drive cost-push inflation. This can erode purchasing power and lead to second-round effects if wages or expectations adjust (Blanchard & Riggi, 2013), while limited monetary and fiscal space constrain policy responses.

**Table 1. Asymptotic critical values of the Fourier CUSUM cointegration test**

| $p$ | $k$ | 1% | 5% | 10% |
|---|---|---|---|---|
| $p = 1$ | 1 | 0.3087608 | 0.3453433 | 0.3702899 |
| | 2 | 0.2994366 | 0.3376164 | 0.3602958 |
| | 3 | 0.2983359 | 0.3376654 | 0.3612651 |
| $p = 2$ | 1 | 0.2990014 | 0.3318281 | 0.3556437 |
| | 2 | 0.2934757 | 0.3297304 | 0.3533944 |
| | 3 | 0.2959929 | 0.3331994 | 0.3559223 |
| $p = 3$ | 1 | 0.2893073 | 0.3221990 | 0.3433905 |
| | 2 | 0.2833487 | 0.3136938 | 0.3333621 |
| | 3 | 0.2825690 | 0.3130642 | 0.3330271 |
| $p = 4$ | 1 | 0.2814507 | 0.3131563 | 0.3318605 |
| | 2 | 0.2712004 | 0.3051386 | 0.3224608 |
| | 3 | 0.2700021 | 0.3024737 | 0.3207691 |

Note: The table presents the critical values for the Fourier CUSUM cointegration test. All critical values were derived using 1,000 observations and 10,000 replications.

An asymmetric effect emerges: increases in oil prices exert a stronger inflationary impact than declines, largely due to rigidities such as menu costs and supply chain frictions. Although Tunisia's subsidy system helps buffer long-term inflation, short-term volatility remains, necessitating carefully balanced policy measures.

In contrast, rising gas prices appear to be linked with lower inflation, reflecting the specific structure of Tunisia's energy market and recent domestic developments. The commencement of the Nawara gas field in 2020 reduced the country's energy deficit and import dependence, easing pressure on energy costs and supporting price stability.

Additionally, Tunisia receives in-kind gas royalties through the Transmed pipeline, further enhancing domestic supply at no direct cost. Together with increased local production, this supports the subsidy framework and helps stabilize energy expenses, contributing to disinflation especially during periods of global price spikes.

This inverse relationship may also stem from regulated energy tariffs and the predominance of gas-based electricity generation, both of which help contain inflation in energy-intensive sectors like manufacturing and agriculture.

Finally, geopolitical uncertainty shows no significant effect on inflation, likely due to a combination of exchange rate controls, regulated pricing, limited exposure of financial markets, and robust institutional buffers that absorb external shocks.

## VI. Conclusion

The findings show no long-term relationship between inflation and oil prices in Tunisia. Temporary inflationary effects from increased oil prices are mostly offset by subsidies on hydrocarbons and electricity. These subsidies impact public finances while supporting households and the broader economy. A gradual reduction of subsidies could redirect resources toward other investments, encourage energy efficiency, and facilitate renewable energy initiatives. Such changes may require social support measures to assist vulnerable groups.

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**Table 2. Empirical size**

| | | $k=1$ | | | | $k=2$ | | | | $k=3$ | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| $p$ | $\alpha$ | $n=100$ | $n=200$ | $n=300$ | $n=500$ | $n=100$ | $n=200$ | $n=300$ | $n=500$ | $n=100$ | $n=200$ | $n=300$ | $n=500$ |
| 1 | 0 | 0.1092 | 0.083 | 0.0756 | 0.0598 | 0.0918 | 0.0660 | 0.0584 | 0.0452 | 0.11 | 0.0808 | 0.08 | 0.0628 |
| | 0.2 | 0.0242 | 0.0116 | 0.0088 | 0.006 | 0.0272 | 0.0146 | 0.01 | 0.0056 | 0.0300 | 0.0166 | 0.0114 | 0.0064 |
| 2 | 0 | 0.1112 | 0.0820 | 0.0666 | 0.0540 | 0.1430 | 0.1044 | 0.0868 | 0.0734 | 0.1516 | 0.1146 | 0.0976 | 0.0822 |
| | 0.2 | 0.0248 | 0.0116 | 0.0072 | 0.0042 | 0.0334 | 0.0132 | 0.0094 | 0.0076 | 0.0390 | 0.0174 | 0.0124 | 0.0096 |
| 3 | 0 | 0.1122 | 0.0798 | 0.0742 | 0.0610 | 0.1168 | 0.0840 | 0.0740 | 0.0580 | 0.1142 | 0.0826 | 0.0722 | 0.0574 |
| | 0.2 | 0.0238 | 0.0104 | 0.0064 | 0.0036 | 0.0262 | 0.0072 | 0.0058 | 0.0044 | 0.0258 | 0.0072 | 0.0056 | 0.0044 |
| 4 | 0 | 0.1222 | 0.0778 | 0.0782 | 0.0606 | 0.1256 | 0.0816 | 0.0786 | 0.0588 | 0.1196 | 0.0838 | 0.0752 | 0.0592 |
| | 0.2 | 0.0270 | 0.0092 | 0.0064 | 0.0036 | 0.0254 | 0.0088 | 0.0058 | 0.0026 | 0.0272 | 0.0108 | 0.0048 | 0.0032 |

Notes: The table presents the empirical size of the Fourier cointegration CUSUM test. All values are based on 5,000 replications. ($\gamma$) denotes the autoregressive coefficient, (p) the number of regressors, (k) the Fourier frequency, and (n) the number of observations.

**Table 3. Empirical power**

| | k/n | 100 | 200 | 300 | 500 |
|---|---|---|---|---|---|
| $p=1$ | 1 | 0.8940 | 0.9186 | 0.9260 | 0.9414 |
| | 2 | 0.8658 | 0.8926 | 0.9012 | 0.9166 |
| | 3 | 0.8920 | 0.9206 | 0.9208 | 0.9354 |
| $p=2$ | 1 | 0.8708 | 0.904 | 0.9158 | 0.9324 |
| | 2 | 0.8602 | 0.8974 | 0.913 | 0.9276 |
| | 3 | 0.852 | 0.884 | 0.9012 | 0.9172 |
| $p=3$ | 1 | 0.8884 | 0.9194 | 0.9262 | 0.9402 |
| | 2 | 0.8822 | 0.9078 | 0.928 | 0.928 |
| | 3 | 0.8804 | 0.9054 | 0.9254 | 0.9372 |
| $p=4$ | 1 | 0.8748 | 0.9046 | 0.9252 | 0.9378 |
| | 2 | 0.8736 | 0.9062 | 0.9278 | 0.9332 |
| | 3 | 0.8756 | 0.9092 | 0.9302 | 0.9376 |

Notes: The table presents the empirical power results for the Fourier CUSUM cointegration test. All values are computed based on 5000 replications. (p) refers to the number of regressors, (k) denotes the Fourier frequency, and (n) indicates the number of observations.

**Table 4. Estimation results**

| | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | **Panel A: Unit Root test** | | | | | | |
| | **Fourier ADF** | **ADF** | **F-test** | **Zivot-Andrews** | | | | | | | |
| Series | Level | Lags | K | $1^{st}$ diff | lags | K | Level | $1^{st}$ diff | | t-statistic | Break point |
| *inf* | -1.263 | 4 | 0.82 | -12.142*** | 3 | 0.92 | 0.472 | -13.371*** | 22.9 | -2.883 | 01/01/2000 |
| *wti* | -4.038*** | 2 | 0.76 | -14.769*** | 1 | 1.48 | -2.040 | -14.738*** | 6.930 | -4.143 | 01/06/2014 |
| *gasp* | -4.735*** | 1 | 1.01 | -16.02*** | 1 | 1.2 | -3.093** | -16.015*** | 6.732** | -4.573* | 01/02/1999 |
| *gpr* | -6.234 *** | 2 | 3.66 | -13.010*** | 4 | 3.67 | -7.001 *** | -21.470*** | 4.457*** | -8.946*** | 01/07/2001 |
| | | | | | **Panel B: Cointegration test** | | | | | | |
| | | | | | | | $\mathbf{CS_{476}}$ | | | $\mathbf{CI_f^{m*}}$ | |
| **statistics** | | | | | | | 2.931 | | | 3.348 | |
| | | | | | **Panel C: QARDL Results** | | | | | | |
| $\tau$ | **0.1** | **0.2** | **0.3** | **0.4** | **0.5** | **0.6** | **0.7** | **0.8** | **0.9** | | |
| **Short-run** | | | | | | | | | | | |
| $\phi_{inf_{t-1}}$ | 1.1892*** | 1.2009*** | 1.2456*** | 1.2577*** | 1.2723*** | 1.2555*** | 1.2536*** | 1.3408*** | 1.4473*** | | |
| $\phi_{inf_{t-2}}$ | -0.1891*** | -0.2007*** | -0.245 | -0.25783*** | -0.27236*** | -0.25561*** | -0.25406*** | -0.34133*** | -0.44840*** | | |
| $\gamma_{wti}$ | 0.0007 | 0.001** | 0.00130*** | 0.00128** | 0.00139*** | 0.00140*** | 0.00169** | 0.00103* | 0.00150 | | |
| $\gamma_{gasp}$ | -0.001* | -0.001*** | -0.00122*** | -0.00110** | -0.00125*** | -0.00158*** | -0.00172** | -0.00171*** | -0.00254** | | |
| $\gamma_{gpr}$ | 0.0007 | 0.00014 | -0.00007 | -0.00009 | 0.00060 | 0.00075 | 0.00090 | 0.00108* | 0.00054 | | |
| **Long-run** | | | | | | | | | | | |
| $\beta^*_{wti}$ | -11.827 | -8.61942 | 10.24419 | 15.1348 | -49.5177 | 29.0676 | 3.9640 | 2.1716 | 1.36794 | | |
| $\beta^*_{gasp}$ | 20.656 | 10.09245 | -9.58743 | -12.9617 | 44.7210 | -32.7172 | -4.0542 | -3.6055 | -2.31329 | | |
| $\beta^*_{gpr}$ | -12.701 | -0.88879 | -0.55185 | -1.1228 | -21.4817 | 15.4345 | 2.1073 | 2.2857 | 0.49228 | | |

Note: This table presents the unit root tests in Panel A, where the optimal k frequency is determined using the AIC criterion. Panel B provides the cointegration test results following Tsong et al. (2016), with critical values of 0.098, 0.132, and 0.215 corresponding to 1%, 5%, and 10% significance levels, respectively. Panel C displays the QARDL results. * indicates p<0.01, indicates p<0.05, and * indicates p<0.10.

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