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Middle Eastern Oil Shares and Brent Prices: A Composite Index and Exploratory ARDL Analysis

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ABSTRACT

This study proposes the Middle East Oil Influence Index (MEOII), an equally weighted geometric mean of the region's shares of world oil reserves, production and crude oil exports. The index describes a relative market position; it does not directly measure market power or geopolitical influence. An exploratory autoregressive distributed lag (ARDL) analysis relates the index to the natural logarithm of annual average Brent prices, with the Index of Global Real Economic Activity (IGREA) as a demand-related control. The stated period is 2000-2025. The reported specification is ARDL(1,2,0), with a bounds F-statistic of 5.607 and an adjustment coefficient of -0.264. Positive short-run associations are reported, whereas neither explanatory variable has a reported statistically significant long-run coefficient. Incomplete documentation of the component series, variable transformation, unit-root results and bounds-test settings limits inference. The estimates therefore remain exploratory and do not establish cointegration, a causal price effect or the absence of a long-run relationship. The contribution is a transparent index formula and a framework for investigating regional oil shares alongside global demand conditions.

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1. Introduction

Oil price changes affect production costs, fiscal revenues and energy expenditure. Their interpretation depends on the shocks behind them. Supply disruptions, changes in demand and expectations can produce different economic responses (Kilian, 2008, 2009; Baumeister & Kilian, 2016a). An empirical association with prices therefore needs a more limited interpretation than a causal supply effect.

The Middle East is a major oil-producing and exporting region. Its resource base, current output and exports describe different aspects of its participation in oil markets (Energy Institute, 2026). Reserves concern resources recoverable under specified conditions; production and exports concern flows. None of these measures alone establishes the ability of producers to set prices.

This study combines the region's shares of reserves, production and crude oil exports in the Middle East Oil Influence Index (MEOII). The name identifies the proposed indicator. Its operational meaning is a composite of relative shares, rather than an independently validated measure of influence.

We examine how the index relates to Brent prices in an annual ARDL specification with a demand-related control. The construction is intended to make the three components explicit. It does not establish that a composite is superior to separate indicators, since the study does not compare their explanatory or forecasting performance.

The annual sample is stated as 2000-2025. The demand-related control is the Index of Global Real Economic Activity, associated with Kilian (2009, 2019). Including this variable accounts for one observed dimension of demand conditions; it does not isolate an exogenous regional supply shock.

The reported estimates suggest positive short-run associations and no individually significant long-run coefficients for the two explanatory variables. These summaries are interpreted cautiously because the full coefficient and test results are not documented. In particular, a reported bounds statistic without its testing case and reference values cannot establish a verified cointegration conclusion.

The study's main contribution is the proposed index construction. The empirical exercise illustrates questions that this index can address, subject to data comparability, a short annual sample and incomplete estimation documentation.

2. Background and related research

2.1. *Oil prices and regional market position*

Kilian (2009) distinguishes supply shocks, demand for industrial commodities and oil-specific demand shocks. This distinction matters here: oil production and prices may respond to the same demand conditions, and producers may adjust output in response to prices. An observed relationship between regional shares and prices need not identify the direction of causation.

Baumeister and Kilian (2016a) review the causes of major oil price movements and the difficulty of predicting them. Their analysis of the decline beginning in June 2014 also considers demand, production and expectations (Baumeister & Kilian, 2016b). These studies do not support attributing a price episode to one regional indicator without identifying the relevant shocks.

Kaufmann et al. (2004) examine OPEC and oil prices. OPEC membership and the Middle East are different classifications: some OPEC producers are outside the region, while some regional producers are outside OPEC. Evidence about OPEC cannot be treated as a direct estimate of a Middle Eastern regional effect.

Inventories and expectations are further dimensions of oil price formation (Kilian & Murphy, 2014). Geopolitical risk is another distinct concept for which Caldara and Iacoviello (2022) develop a news-based measure. MEOII does not include inventories, production flexibility, transport vulnerabilities or geopolitical risk. It should not substitute for these measures.

The index combines resource endowment, current output and international sales. These components can be related, and production and exports may overlap economically. Equal weighting does not eliminate correlated information or establish the independent contribution of each component.

2.2. *Measuring demand and composite indicators*

IGREA is a shipping-based measure of cyclical activity in industrial commodity markets, rather than a direct measure of world GDP or oil consumption. The Dallas Fed describes the updated series as percentage deviations from trend and incorporates the correction discussed by Kilian (2019) (Federal Reserve Bank of Dallas, n.d.). Hamilton (2021) questions aspects of shipping-based measurement and favours world industrial production as an alternative. The choice of demand proxy is therefore a substantive modelling decision.

The OECD, European Union and EC-JRC (2008) discuss the importance of theory, weighting, aggregation and sensitivity analysis when constructing composite indicators. Combining variables produces a summary whose interpretation depends on those choices. The present study proposes a particular summary of oil shares, without claiming an exhaustive review or priority over every previous composite measure.

The empirical question is whether changes in this summary are associated with Brent prices after conditioning on IGREA and the specified dynamics. A comparison with separate component models and alternative demand proxies would be needed to evaluate whether the index adds explanatory value.

3. Construction of the Middle East Oil Influence Index

3.1. *Components and scope*

For year t, let $R_{\mathrm{ME},t}$ and $R_{\mathrm{W},t}$ denote regional and world proved oil reserves; $Q_{\mathrm{ME},t}$ and $Q_{\mathrm{W},t}$ denote regional and world oil production; and $X_{\mathrm{ME},t}$ and $X_{\mathrm{W},t}$ denote regional and world crude oil exports. Within each ratio, numerator and denominator must refer to the same year, product coverage and measurement unit.

The three relative shares are defined as follows:

$$\mathrm{ORS}_t = R_{\mathrm{ME},t} / R_{\mathrm{W},t}$$

$$\mathrm{OPS}_t = Q_{\mathrm{ME},t} / Q_{\mathrm{W},t}$$

$$\mathrm{OES}_t = X_{\mathrm{ME},t} / X_{\mathrm{W},t}$$

ORS is the reserve share, OPS the production share and OES the crude oil export share. A regional aggregate requires a consistent set of countries across years and components. The estimation documentation does not list that set or the component series identifiers, which limits the reproducibility of the reported index.

3.2. *Aggregation and interpretation*

$$\mathrm{MEOII}_t = (\mathrm{ORS}_t \times \mathrm{OPS}_t \times \mathrm{OES}_t)^{1/3}$$

For positive fractional shares, the formula yields a dimensionless value between zero and one. Each component has an exponent of one third: a given proportional increase in any one component has the same proportional contribution to the index, holding the others fixed. A zero component makes the product zero; missing observations do not justify assigning zero.

A higher index value indicates a higher combined relative share. It does not necessarily imply an increase in absolute output or exports, because a regional share can rise when the rest of the world's activity falls. The formula gives no direct estimate of pricing power.

3.3. *Rationale and limitations of the geometric mean*

Geometric aggregation responds to proportional changes and gives a lower score to an imbalanced set of components than its arithmetic mean. It permits trade-offs between positive components, although compensation differs from additive aggregation (OECD et al., 2008). It does not guarantee robustness to outliers or prevent one component from driving changes over time.

Equal exponents are a modelling choice. They do not establish equal empirical importance or remove overlapping information. The study does not report comparisons with alternative weights or aggregation rules, so superiority over those alternatives is not established.

Expressing all three shares as percentages instead of fractions multiplies the composite by 100. That change affects the numerical interpretation of a level coefficient. The empirical index scale is not documented, so the reported summaries cannot support a numerical effect per index point.

3.4. *Research question*

The expected association with prices is ambiguous. Increased regional output or exports may add to available supply, while a rising regional share can also reflect disruption elsewhere or changing expectations. Reserves, production and exports need not move together.

The research hypothesis is that the short-run coefficients on changes in MEOII are jointly different from zero. This is an association hypothesis within the specified model. It is distinct from a test of an individual coefficient and from the bounds test on a levels relationship.

4. Data and variables

4.1. *Sources and coverage*

The stated sample contains 26 annual observations from 2000 through 2025 before lagging or other losses. Annual frequency limits the analysis of rapid price responses and the number of dynamic parameters that can be estimated.

The study identifies the Energy Institute's Statistical Review of World Energy 2026 as the source of its oil series. The edition is a documented publication, but citing the edition does not identify the exact reserve, production and export series used in estimation. Table identifiers, the regional country list, the data vintage and the treatment of missing values are not specified.

Comparable product definitions also matter. For example, the Energy Institute's 2025 methodology includes condensates and natural gas liquids in its broad oil production definition and notes that its reserves tables were not updated in that edition (Energy Institute, 2025). A broad oil series cannot be relabelled as crude oil alone. These documentation issues prevent treating the index as a verified, consistently defined annual crude-only series through 2025.

The stated Brent variable is the annual average price in current U.S. dollars per barrel. No price deflator is specified, so the outcome is interpreted as nominal. The stated IGREA aggregation is the arithmetic average of the twelve monthly values for each year. The exact downloaded version used in estimation is not identified; access to an updated series does not establish which vintage generated the reported estimates.

4.2. *Transformations and coefficient interpretation*

Let $y_t = \ln(P_t)$, where P_t is the annual average Brent price. The logarithm of the annual mean is not the mean of monthly log prices. Only the former is specified here. Taking logs does not by itself remove heteroskedasticity, trends or structural breaks.

The model equations specify MEOII in levels, and that convention is used below. The study's numerical results are not accompanied by the estimation output needed to verify the implemented transformation. They therefore do not support an elasticity interpretation or a quantified effect per unit of the index.

IGREA remains in levels because it can take negative as well as positive values. Its units describe percentage deviations from trend; a one-point increase is not a one-percent increase in global output. In a log-price equation with level regressors, a coefficient describes a change in log price per regressor unit. For a finite change d and coefficient b, the implied conditional percentage price change is $100 \times [\exp(bd) - 1]$, rather than an elasticity.

4.3. *Variable definitions*

4.4. *Scope of the controls*

MEOII summarises regional oil shares, while IGREA represents one demand-related dimension. Neither variable is a pure structural shock. Their inclusion does not control for every supply or demand determinant, and the model does not separate feedback between prices and production decisions.

Table 1 Variables in the specified model

Variable	Definition and scale	Role	Expected sign
y or LnOIL	ln of annual average nominal Brent price in USD per barrel; log measure	Dependent variable	Not applicable
m or MEOII	Geometric mean of three regional shares; levels in the stated equation; implemented scale and transformation not verified	Main explanatory variable	Ambiguous
z or IGREA	Annual arithmetic mean of monthly activity index; percentage deviations from trend; no log	Demand-related control	Positive hypothesised

A positive association with IGREA is plausible, but it remains an empirical question. Exchange rates, inventories, interest rates, non-regional supply and geopolitical conditions may also affect prices. The short sample constrains the number of controls and does not remove omitted-variable bias.

5. Econometric framework

5.1. *Dynamic specification*

The ARDL framework allows lagged prices and explanatory variables in one equation. Bounds inference requires an appropriate integration structure, deterministic specification and reference distribution, including the absence of I(2) variables (Pesaran et al., 2001). A short sample alone does not guarantee that the approach gives reliable inference.

$$y_t = c + \Sigma_{i=1}^{p}\, a_i\, y_{t-i} + \Sigma_{j=0}^{q_m}\, b_j\, m_{t-j} + \Sigma_{k=0}^{q_z}\, g_k\, z_{t-k} + e_t \quad (1)$$

Here m_t denotes MEOII in levels, z_t denotes IGREA in levels, c is an intercept and e_t is the regression disturbance. The dependent-variable sum starts at lag one; the explanatory-variable sums include the contemporaneous terms.

The reported choice is ARDL(1,2,0), with the variables ordered as y, m and z. Thus the model contains y_{t-1}, m_t, m_{t-1}, m_{t-2} and z_t. Zero is the maximum lag on z; it does not mean that the contemporaneous variable is absent.

5.2. *Integration and breaks*

The stated testing procedure uses augmented Dickey-Fuller tests (Dickey & Fuller, 1981) and a test allowing an unknown break in the trend function, associated with Vogelsang and Perron (1998). A reproducible classification requires the test statistic, deterministic terms, lag rule, break specification and critical values for each variable.

Stationarity around a broken deterministic trend differs from stationarity after first differencing. These outcomes cannot be collapsed into an unspecified I(0)/I(1) classification. Break detection in one unit-root equation also does not show that the estimated price relationship is stable across that break.

The reported summary identifies a 2015 break for log Brent prices, but does not document the complete level and first-difference results for every series. Accordingly, exclusion of I(2) behaviour and the relevant deterministic specification remain unverified.

5.3. *Lag selection and sample size*

The study reports selection by the Akaike information criterion (Akaike, 1974). This compares the candidate models considered; it does not establish a uniquely correct lag structure. Maximum candidate lags and the common selection sample are not documented.

For ARDL(1,2,0), the levels equation is:

$$y_t = c + a\, y_{t-1} + b_0\, m_t + b_1\, m_{t-1} + b_2\, m_{t-2} + g\, z_t + e_t \quad (2)$$

There are five slope coefficients and an intercept. Two lags would leave at most 24 estimation observations from a complete 26-year dataset without pre-sample observations. Missing values or a longer common selection window could reduce that number further; the effective estimation sample is not reported.

5.4. *Levels relationship and error correction*

A levels test concerns a joint restriction. Its decision depends on the deterministic case, the number of explanatory variables and the applicable critical values. An F-statistic above the relevant upper bound can support rejection; a value between the bounds is inconclusive. The statistic alone is not a sufficient decision rule.

The reported bounds statistic is 5.607. The testing case, critical-value table and effective sample are not documented, nor is any t-bounds assessment where applicable. A definitive rejection at the 5% level is therefore not asserted here. Small-sample calibration and the integration and break assumptions also affect its interpretation.

5.5. *Equivalent dynamic representation*

An algebraic rewriting of Equation (2), with Δ denoting a first difference, gives:

$$\Delta y_t = c + \lambda\, y_{t-1} + \delta_m\, m_{t-1} + g\, z_{t-1} + b_0\, \Delta m_t - b_2\, \Delta m_{t-1} + g\, \Delta z_t + e_t \quad (3)$$

The coefficient identities are λ = a - 1 and $\delta_m = b_0 + b_1 + b_2$. Because the selected lag for z is zero, its lagged-level and first-difference coefficients in this rewriting are the same g. They are not independently estimated short-run and level parameters. Equation (3) states this restriction; it does not introduce a new regression result.

If a meaningful levels relation exists and 1 - a is nonzero, the implied coefficients are θ_0 = c / (1 - a), $\theta_m = (b_0 + b_1 + b_2) / (1 - a)$ and θ_z = g / (1 - a). This relationship does not by itself establish cointegration.

A negative adjustment coefficient is consistent with convergence in a suitably specified conditional model. It does not prove a stable economic equilibrium. Where a t-bounds test is applicable, its reference distribution is non-standard. A p-value without an identified reference distribution is insufficient for this decision.

5.6. *Diagnostics and inferential limits*

The stated diagnostic procedure includes Breusch-Godfrey tests for residual serial correlation, Breusch-Pagan-Godfrey tests for heteroskedasticity, Jarque-Bera residual-normality tests, Ramsey RESET and CUSUM/CUSUMSQ stability checks.

Failure to reject a diagnostic null is limited evidence against the particular alternative tested. It does not prove homoskedasticity, normality, correct specification or parameter stability, especially in a short sample. The Durbin-Watson statistic is not an adequate stand-alone check when the equation contains a lagged dependent variable.

6. Reported empirical findings

6.1. *Numerical summaries*

Table 2 brings together the numerical results reported in the analysis. These are estimation summaries, not a reconstruction of the annual data or a re-estimated model. Means, dispersion, skewness and kurtosis are not documented numerically, so no distributional conclusion about the variables is drawn from them.

Table 2 Reported estimation summaries

Item	Reported value	Interpretive limit
Stated period	2000-2025	26 years before lags; effective sample unspecified
Selected model	ARDL(1,2,0)	Variable order y, MEOII, IGREA; AIC search settings unspecified
R^2	0.796	Fit for log prices in the estimation sample
Overall regression F	14.01; p < .01	Different null from the bounds test; degrees of freedom unspecified
Durbin-Watson	1.918	Not a sufficient serial-correlation check with lagged prices
Bounds F	5.607	Case and critical values unspecified; no verified rejection
Adjustment coefficient	-0.264; reported p < .01	Conditional interpretation; test and reference distribution unspecified
Price break year	2015	Reported endogenous break; full test output unspecified

The unit-root account describes stationarity after a break but does not provide enough detail to assign a verified integration order to each variable. The 2015 log-price break is retained as a reported finding, without inferring that it identifies the cause of the price decline.

6.2. *Fit and coefficient interpretation*

The reported R^2 of 0.796 describes fit to log Brent prices within the estimation sample. It is not a causal attribution to MEOII, an out-of-sample forecasting measure or a proportion of variation explained on the original dollar-price scale. Lagged prices are themselves part of the fitted model.

The reported overall regression F-statistic is 14.01, with p < .01 as stated in the analysis. This differs from the bounds F-statistic of 5.607: the two concern different hypotheses. Without the full output, exact degrees of freedom and test settings cannot be checked.

The coefficient narrative describes the contemporaneous and lagged MEOII level terms as individually insignificant, yet also describes a positive significant short-run association. An individual term, a transformed coefficient and a joint test of short-run terms are different hypotheses. The coefficient vector, standard errors and joint test are not reported, so the claimed short-run significance is not treated as established.

6.3. *Levels test and adjustment*

The bounds statistic of 5.607 is retained without declaring verified cointegration. The missing reference values and testing case matter for the decision, while the short sample increases uncertainty. A significant joint levels

test would not require every individual long-run coefficient to be significant.

The reported adjustment coefficient is -0.264. Conditional on a valid error-correction representation, it corresponds to correction of about 26.4% of the preceding period's equilibrium error per year, holding other dynamics fixed. It does not mean that every price shock disappears at that rate. The accompanying $p < .01$ does not identify its test or reference distribution. It therefore cannot establish that an applicable bounds-testing requirement has been met.

The analysis describes positive short-run associations for MEOII and IGREA. Their sizes and uncertainty are not reported numerically. For IGREA, the zero-lag specification also constrains its dynamic representation as shown in Equation (3). A separate, freely estimated short-run IGREA effect is therefore not inferred.

6.4. *Diagnostics and long-run coefficients*

The analysis states that its residual and stability checks did not reject the relevant null hypotheses. Test statistics, p-values, lag choices and the CUSUM/CUSUMSQ plots are not documented, so these statements do not establish that the model satisfies all assumptions. The reported Durbin-Watson value of 1.918 does not resolve this limitation.

Neither MEOII nor IGREA is reported to have an individually significant long-run coefficient. The coefficient estimates, standard errors and confidence intervals are not given. This is insufficient evidence of an independent long-run association within the reported specification; it is not evidence that the corresponding economic effect is zero.

Nor does an insignificant long-run coefficient establish that short-run effects are stronger. Comparing horizons requires the relevant estimates and a test of their difference. The paper does not document such a comparison.

7. Discussion

The proposed index gives an explicit summary of three regional shares. Its usefulness depends on whether the component definitions and annual coverage are comparable. It does not measure spare production capacity, coordination among producers or the ability to change prices. Such influence cannot be inferred from the index label.

Changes in non-regional production, technologies or inventories could affect the relationship between regional shares and prices. This model does not identify those mechanisms. The absence of individually significant long-run coefficients cannot establish that diversification has reduced regional influence.

Positive short-run coefficients, if supported by a fully documented estimation, would describe conditional associations in annual data. They would not identify immediate trading responses or causal effects of production policy. Simultaneity and unobserved common shocks remain plausible explanations.

The reported negative adjustment coefficient is compatible with one interpretation of the fitted dynamics. Its economic meaning depends on a valid levels relation and stable specification. It therefore cannot establish that oil markets return to a uniquely determined equilibrium after every disruption.

The use of IGREA recognises the relevance of commodity-market demand conditions. It does not verify a particular demand-transmission channel in this sample. Alternative activity measures are relevant because the literature debates the construction and interpretation of shipping-based indices (Kilian, 2019; Hamilton, 2021).

The exploratory exercise can inform a more reproducible empirical comparison. That comparison would identify the regional coverage and annual component data, distinguish crude oil from broader liquids, disclose transformations and estimate alternative component and composite specifications on the same sample. These are extensions of the analysis, rather than results demonstrated here.

8. Conclusion and limitations

This study proposes MEOII as the geometric mean of three Middle Eastern shares in world oil markets. It combines resource endowment, current production and crude oil exports, with equal exponents. The result is a descriptive relative-share index, whose interpretation depends on source definitions and weighting choices.

The stated annual ARDL exercise covers 2000-2025 and reports a specification of (1,2,0), a bounds F-statistic of 5.607 and an adjustment coefficient of -0.264. The qualitative summaries describe positive short-run associations and individually insignificant long-run coefficients. Incomplete data and estimation documentation prevents treating these statements as verified significance, cointegration or causal findings.

The index may help organise information used to monitor regional oil participation alongside demand conditions. The analysis does not evaluate the effect of a production agreement, an investment policy or a diversification programme. It therefore provides no quantitative basis for prescribing such interventions or promising reduced volatility.

Several limitations affect interpretation. The sample is short and annual, and the effective estimation period is not specified. The regional country list, component series, scale, reserve coverage and missing-value treatment are not documented in sufficient detail. The implemented MEOII transformation is uncertain. Unit-root and break specifications, bounds reference values, coefficient uncertainty and diagnostic output are incomplete. The model also omits other price determinants and does not identify exogenous shocks.

Further research can examine the index with documented annual components and sensitivity to weights and aggregation, compare it with separate-share models and alternative demand proxies, and assess finite-sample and break-sensitive inference. Until those comparisons are

available, the index is a proposed descriptive tool and the reported price relationships remain exploratory.

Declarations

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