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Food import demand with structural breaks, economic embargo and the COVID-19 pandemic in a wealthy, highly import-dependent country

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Abstract

Purpose – Qatar, a wealthy country with an open economy has limited arable land. To meet its domestic food demand, the country heavily relies on food imports. Additionally, the over three year-long economic embargo enforced by regional neighbors and the covariate shock of the COVID-19 pandemic have demonstrated the country's vulnerability to food insecurity and potential for structural breaks in macroeconomic data. The purpose of this paper is to examine short- and long-run determinants of Qatar's imports of aggregate food, meats, dairy and cereals in the presence of structural breaks.

Design/methodology/approach – The authors use 24 years of food imports, gross domestic product (GDP) and consumer price index (CPI) data obtained from Qatar's Planning and Statistics Authority. They use the autoregressive distributed lag (ARDL) cointegration framework and Chambers and Pope's exact nonlinear aggregation approach.

Findings – Unit root tests in the presence of structural breaks reveal a mixture of I (1) and I (0) variables for which standard cointegration techniques do not apply. The authors found evidence of a significant long-run relationship between structural changes and food imports in Qatar. Impulse response functions indicate full adjustments within three-quarters of a year in the event of an exogenous shock to imports.

Research limitations/implications – An exogenous shock of one standard deviation on this variable would reduce Qatar's food imports by about 2.5% during the first period but recover after the third period.

Originality/value – The failure of past aggregate food demand studies to go beyond standard unit root testing creates considerable doubt about the accuracy of their elasticity estimates. The authors avoid that to provide more credible findings.

Keywords Food import demand, Structural break, Bounds test, Autoregressive distributed lag (ARDL), Cointegration, Qatar

Paper type Research paper

1. Introduction

Qatar's economy is built primarily on high levels of oil and gas production and export. Coupled with the country's relatively small size in terms of land area and population, this explains why Qatar is one of the richest countries on a per capita basis. Although wealthy, Qatar has a limited amount of arable land and consequently produces very little food. It is thus a highly

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import-dependent country, with a food import profile that involves trade relationships with more than 100 countries annually (Kaitibie *et al.*, 2017). Many of the imported food types (e.g. meats), are highly concentrated at the country of origin (Basher *et al.*, 2013). Understanding the food import situation is crucial for the successful implementation of the country's well-developed food security plan. The country's food security plan has two primary aims: to improve domestic food production and streamline international trade. The country's heavy reliance on food imports exposes it to vulnerabilities resulting from shocks in supply, as was experienced during the 2017–2021 economic embargo and the recent breakdown in international food supply logistics due to the COVID-19 pandemic (Kaitibie *et al.*, 2022).

While aggregate food demand analysis has a long history, it remains a fertile research area for generating policy-relevant information and for testing the economic theory (Boysen, 2016; Hoang, 2018; Bairagi *et al.*, 2020). Since Sims' (1980) study, most aggregate time series food demand analyses have used vector autoregression (VAR) methods. The most commonly-used methods to assess the existence of long-run relationships are Engle and Granger's (1987) two-step residual-based cointegration procedure and Johansen's (1991) system-based reduced-rank approach. A prerequisite for using these tests is that the underlying variables should follow a random walk process (Noriega and Ventosa-Santaularia, 2012). Many methods exist for testing the unit root hypothesis in VARs. These methods include the Dickey–Fuller (DF) (Dickey and Fuller, 1979), the augmented Dickey–Fuller (ADF) (Dickey and Fuller, 1981), the DF generalized least squares (DF-GLS) (Elliott *et al.*, 1996), the Phillips–Perron (PP) (Phillips and Perron, 1988), the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (Kwiatkowski *et al.*, 1992) and the Leybourne ADFmax (Leybourne, 1995) tests. A key characteristic of these tests is that they generally ignore the possible occurrence of regime change in the series' intercept, slope or both, potentially leading to erroneous classification of the order of integration.

Gregory and Hansen (1996) observed that the presence of regime change in time series data diminishes the power of traditional unit root tests in detecting stationarity. Further, Perron (1989) showed that the power of standard tests to reject a unit root in the presence of a neglected structural break decreases when the stationary alternative is true. Under such circumstances, traditional cointegration techniques (e.g. Engle and Granger's test and Johansen's test) and their variants are rendered invalid. In cases such as these, researchers should use autoregressive distributed lag (ARDL) models (Pesaran and Shin, 1999). The ARDL produces consistent estimates and, in studies with small samples and single-equation settings, is superior to Phillips and Hansen's (1990) semi-parametric cointegration approach. Furthermore, the ARDL procedure works well in models with a mixture of purely I (1), purely I (0) or mutually cointegrated endogenous regressors without the need for prior testing for the order of integration.

This study demonstrates the utility of the ARDL framework in circumstances where there is uncertainty regarding the order of integration of the underlying variables. We consider Qatar's import demand for food. Specifically, we estimate both the long-run prices, income and effective exchange rate elasticities of Qatar's aggregate food import demand and the dynamics of short-run adjustments of food imports to changes in these variables. Previous food demand studies only employ traditional unit root tests (see for example, Niemi, 2004; Nguyen and Jolly, 2013); they extracted long- and short-run elasticities from the cointegration vector and the error correction model (ECM), regardless of the diminished sensitivity of these tests to aberrations in the time series. Perron (1989) has argued that most macroeconomic variables are trend stationary if one allows a single break point in the intercept and slope. The failure of past aggregate food demand studies to go beyond standard unit root testing creates considerable doubt about the accuracy of their elasticity estimates. The generation of accurate price and income elasticities is not only crucial in the analysis of food consumption behavior, but is also of particular interest to policymakers as it enables them to design effective price and income support policies and programs (Nzuma and Sarker, 2010).

Given the foregoing, we build on the previous literature by estimating aggregate food import demands in the presence of structural breaks using the exact nonlinear aggregation approach of Chambers and Pope (1992). This approach derives market demand from aggregate budget shares of expenditures of multiple consumers. It led to the well-known price-independent generalized logarithmic (PIGLOG) linearity family of demand functions later popularized by Muellbauer (1975, 1976). We employ time series data from Qatar, covering the period of a regional economic embargo on Qatar and the COVID-19 pandemic. Given Qatar's heavy reliance on food imports, its open economy, its small size in international food trade and related macroeconomic variables are repeatedly subject to international and domestic shocks, rendering them particularly amenable to exact nonlinear aggregation demand analysis.

2. Literature review

International trade is crucial to the growth and development of an economy. In that regard, several empirical studies have investigated the factors that influence food import demand, using varying methods and scope of coverage. Ho (2004) employed the Johansen–Juselius maximum likelihood cointegration and error correction technique to test aggregate and disaggregated import demand models using quarterly data for Macao. For the chosen period, the study found that there was long-run cointegration in the disaggregated model and that the signs of the estimated coefficients were inconsistent with the economic theory.

In another study, Aljebrin and Ibrahim (2012) used the panel seemingly unrelated regression (SUR) model to investigate the determinants of import demand for countries in the Gulf Cooperation Council (GCC), including Qatar. Using panel data for the period of 1994–2008, they found that real income, international reserves, private consumption and gross capital formation positively affected import demand, both in the short- and long-run. Amiri and Talbi (2012) also estimated the import demand function of oil-exporting countries. Using the panel cointegration technique, they found that import demand was positively influenced by the real exchange rate, oil prices and domestic demand.

Another related study by Hoang (2018) analyzed food demand and the short-term impacts of potential market shocks on quantity and calorie consumption in Vietnam. The study's findings indicated inelastic own price and expenditure elasticities compared to other foods. As a result, Hoang (2018) recommended that the government should provide necessary safety net programs for the poor. Relatedly, Bairagi *et al.* (2020) estimated a demand system of 15 major food items in Vietnam. They found a large variation in the estimated price and expenditure elasticities. Bairagi *et al.* (2020) recommended that government policy should encourage demand-oriented food production, emanating from urbanization and income growth and that farmers should diversify their crops to meet the rising demand for these food products.

Mehmood *et al.* (2013) employed ARDL bounds tests to examine the price and income elasticities of the disaggregated import demand function spanning the period of 1972–2009. They found evidence of a long-run relationship and real GDP had a positive effect on import demand for food items. Ibrahim (2015) also employed the ECM to assess the long- and short-run determinants of merchandise imports in Saudi Arabia using annual time series data spanning the period of 1975–2011. The study found that in the long- and short-run, real GDP, investment expenditure, government consumption expenditure, and private consumption expenditure had positive effects on the import of merchandise. Using the dynamic system GMM methodology, Asaana and Sakyi (2021) examined the drivers of goods and services imports in 32 sub-Saharan African countries for theperiod of 1990–2016. They found that expenditure components, foreign exchange reserves and the relative import prices were the major drivers.

Given the useful insights advanced by previous studies (i.e. Ho, 2004; Aljebrin and Ibrahim, 2012; Amiri and Talbi, 2012; Mehmood *et al.*, 2013; Ibrahim, 2015; Bairagi *et al.*, 2020; Hoang, 2018; Asaana and Sakyi, 2021; Kaitibie *et al.*, 2022), the lack of similar studies on import-dependent countries warrants further research. Our study fills this gap by investigating similar phenomena and the impacts of both the 2017–2021 economic embargo and the COVID-19 pandemic on Qatar.

3. Methodology

3.1 Data

Data on food imports, GDP and the consumer price index (CPI) were obtained from Qatar's Planning and Statistics Authority (2022) and covered the period of 1998 Q1–2021 Q4. The data contained import values. The quantities of different types of food items (meat, cereals, dairy and an aggregate of all foods) were provided via eight-digit Harmonized System (HS8) codes. The nominal effective exchange rate (NEER) was obtained from the International Monetary Fund (IMF, 2022). Qatar's GDP growth rate was obtained from the World Development Indicators (World Bank, 2022). All the series were log-transformed.

3.2 Analytical framework

To efficiently analyze the short-run and long-run dynamic relationships between the variables mentioned in Section 3.1, the authors employed the ARDL cointegration framework. This framework has been used in several studies (e.g. Kumar *et al.*, 2021; Mustapha and Said, 2016; Akber and Paltasingh, 2020; Chopra, 2022). Prior to estimating the ARDL framework, the authors tested the unit root status of the variables, with and without structural breaks.

3.2.1 Unit roots test. The ARDL technique is only appropriate in situations where some variables are integrated of Order 0 or 1 (i.e. I (0) or I (1)). Thus, it is crucial to conduct unit root tests on all the regressors to ascertain their order of integration. First, we employed the ADFmax unit root test proposed by Leybourne (1995). This test is more powerful to reject a false null hypothesis than the standard DF and ADF tests, but it does not account for structural breaks. Mindful of possible structural breaks in macroeconomic variables due to events like the food price crises of 2008 and 2011, and the economic embargo of 2017–2021, we conducted additional unit root tests in the presence of structural breaks, using the methods of Zivot and Andrews (1992) and Clemente *et al.* (1998). These tests, respectively, allow for single and double structural breaks in both intercept and trend of the data series and estimated optimal break points.

3.2.2 The ARDL cointegration framework. Following Pesaran and Shin (1999), a simple ARDL (p, q) process is expressed as follows:

$$y_t = \beta_0 + \delta t + \sum_{i=1}^{p} \beta_i y_{t-i} + \sum_{j=0}^{q} \alpha_j x_{jt-j} + \varepsilon_t \; \forall t = 1, \; \ldots, \; T \tag{1}$$

where y_t is a $k \times 1$ vector of the dependent variable, x_t is a $k \times k$ matrix of regressors, t is a linear deterministic trend, ε_t is a $k \times 1$ vector of white noise error term with variance σ^2. While β_0, δ, β_i and α_j are $k \times 1$ vectors of unknown parameters, p and q represent lag orders on the dependent and independent variables, respectively. Equation (1) can be rewritten as an ECM:

$$\Delta y_t = \rho_0 + \tau t + \sum_{i=1}^{p} \rho_i \Delta y_{t-i} + \sum_{j=0}^{q} \varpi_j \Delta x_{jt-j} + \varphi v_{t-1} + e_t \tag{2}$$

where Δ is a difference operator, v_{t-1} is the error correction term (ECT) derived from lagged OLS residuals of the cointegrating regression, $y_t = \beta_0 + \delta t + \sum_{i=1}^{k} \alpha_i x_{it} + v_t$, such that $v_{t-1} = y_{t-1} - \widehat{\beta}_0 - \widehat{\delta} t - \sum_{i=1}^{k} \widehat{\alpha}_i x_{it-1}$, e_t is a new error term and φ is the speed of adjustment that measures how fast the system converges to its long-run equilibrium. The short- and long-run parameters are easily extracted from the coefficients of the unrestricted ECM that is expressed as follows:

$$\Delta y_t = \rho_0 + \tau t + \sum_{i=1}^{p} \rho_i \Delta y_{t-i} + \sum_{j=0}^{q} \varpi_j \Delta x_{jt-j} + \varphi \left[\theta_0 y_{t-1} - \widehat{\beta}_0 - \widehat{\delta} t - \sum_{i=1}^{k} \widehat{\alpha}_i x_{it-1} \right] + e_t \tag{3}$$

where the term in square brackets is the ECT. The parameters capturing short- and long-run dynamics are θ_0, $\widehat{\beta}_0$ and $\widehat{\delta}$ and $\widehat{\alpha}_i$, and $\frac{\widehat{\delta}}{\theta_0}$ and $\frac{\widehat{\alpha}_i}{\theta_0}$, respectively.

To incorporate a structural break in the ARDL, the literature suggests two models; hereafter, these models are referred to as Model 1 and 2 (e.g. see Zivot and Andrews, 1992; Narayan and Narayan, 2005). According to Narayan and Narayan (2005), while Model 1 focuses on a change in the intercept, Model 2 allows for a change in both the intercept and the slope. For an AR (k) univariate model, the two models are derived as follows:

$$y_t = \beta_0 + \rho y_{t-1} + \sum_{j=1}^{k} \gamma_j y_{t-j} + \varepsilon_t \tag{4}$$

that can be rewritten as follows:

$$y_t - \rho y_{t-1} - \sum_{j=1}^{k} \gamma_j y_{t-j} = \beta_0 + \varepsilon_t. \tag{5}$$

Hence, assuming a deterministic time trend, t, Model 1 (Narayan and Narayan, 2005, p. 1980) is given as follows:

$$\Delta y_t = \beta_0 + \alpha y_{t-1} + \beta_1 t + \theta_1 dU_t + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + \varepsilon_t \tag{6}$$

while Model 2 is

$$\Delta y_t = \beta_0 + \alpha y_{t-1} + \beta_1 t + \theta_1 dU_t + \pi_1 dT_t + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + \varepsilon_t, \tag{7}$$

where $\alpha = \rho - 1$ and $0 < \rho < 1$. Here, $\rho = 1$ implies that the series is an I (1) process with a random walk with a possible drift. Dummy variables dU_t and dT_t indicate a change in the intercept and in the slope, respectively at Time TB with $dU_t = 1$ and $dT_t = t - TB$ if $t > TB$ and zero otherwise (Narayan and Narayan, 2005). The Zivot and Andrews (1992) test assesses the null hypothesis of a unit root process with drift that excludes structural breaks, against the alternative hypothesis of an estimated structural break in the trend. Lumsdaine and Papell (1997) proposed an extension to Zivot and Andrews' (1992) model that enables researchers to endogenously test for two structural breaks. Models 1 and 2 are, respectively, respecified to take the following forms:

$$\Delta y_t = \beta_0 + \alpha y_{t-1} + \beta_1 t + \theta_1 dU1_t + \omega_1 dU2_t + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + \varepsilon_t \tag{8}$$

and

$$\Delta y_t = \beta_0 + \alpha y_{t-1} + \beta_1 t + \theta_1 dU1_t + \lambda_1 dT1_t + \omega_1 dU2_t + \eta_1 dT2_t + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + \varepsilon_t, \tag{9}$$

where $dU1_t$ and $dU2_t$ are dummy variables, representing a shift in the means occurring at Times TB_1 and TB_2, respectively with $TB_2 > TB_1 + 2$. The terms $dT1$ and $dT2$ are indicators for the shift in trends. As in the one break case, the same procedure was employed to choose the break points. The *t-sig* method was used to select the lag length.

An appropriate unit root test is used to test the null hypothesis of nonstationarity (i.e. $H_0 : \rho - 1 = 0$) allowing for a structural break in both the intercept and the trend against the alternative of stationarity (i.e. $H_A : \rho - 1 \neq 0$). Both models were applied to endogenously determine the breakpoint.

3.3 Model specification

According to Khan and Ross (1977) and Bhatti and Al-Shanfari (2016), and in keeping with Gorman's (1953) exact linear-in-moments aggregation approach, food import demand at Time t, (FM_t), is a function of the import price and the importing country's real income:

$$FM_t = \beta_0 + \beta_1 P_t + \beta_2 RGDP_t + \varepsilon_t, \tag{10}$$

where P_t is the unit price of imported food; $RGDP_t$ is the real GDP; and ε_t is an i.i.d. $\sim N(0, \sigma^2)$ white noise disturbance term. In the absence of domestic prices, Nguyen and Jolly (2013) have suggested using the real food import unit price as a proxy for the domestic price. This figure is obtained by dividing the nominal food import unit value by the CPI. Qatar has a dearth of data on domestic food prices. Given that food import and domestic prices in Qatar do not differ significantly due to low tariffs in the food sector, we used import prices as a proxy for domestic prices. Therefore, Equation (8) becomes the following:

$$FM_t = \beta_0 + \beta_1 \frac{P_t}{CPI_t} + \beta_2 RGDP_t + \varepsilon_t. \tag{11}$$

Both Khan and Ross (1977) and Nguyen and Jolly (2013) have argued that food import demand is essentially a dynamic process that adjusts toward the equilibrium quantity whenever there are any shocks in the market. This dynamicity arises from the fact that imported food takes time to produce, transport and deliver to customers leading to the delayed responsiveness of imports to market changes (Nguyen and Jolly, 2013). Accordingly, Equation (9) can be written as an ARDL process similar to Equation (1):

$$FM_t = \beta_0 + \delta t + \sum_{i=1}^{p} \beta_i FM_{t-i} + \sum_{j=0}^{q} \alpha_j x_{jt-j} + \varepsilon_t \ \forall t = 1, \ \ldots, \ T, \tag{12}$$

where x_{jt} is a set of regressors hypothesized to influence food import demand. As shown in Equation (2), and subject to confirmation of a cointegrating relationship between the dependent variable and the set of regressors, this model can be converted into an ECM by differencing as follows:

$$\Delta FM_t = \beta_0 + \delta t + \sum_{i=1}^{p} \beta_i \Delta FM_{t-i} + \sum_{j=0}^{q} \gamma_j \Delta x_{jt-j} + \varphi v_{t-1} + e_t, \tag{13}$$

where all symbols are as previously defined.

In addition to real import unit price and GDP, previous research (e.g. Sultan, 2011; Towbin and Weber, 2013) incorporates other financial and macroeconomic regressors in the demand equation, including net foreign reserves, net exports, population and the NEER. The current study included the nominal exchange rate (to account for inflationary changes) in x_{jt} because the exchange rate affects the affordability of food imports. We incorporated a single structural break, *SB*, to capture changes in both the intercept and slope, after confirming its presence, because the ARDL does not currently incorporate two or more breakpoints. In keeping with Khan and Ross' (1977) and Boylan *et al.'s* (1980) recommendations, the final food import demand model was specified in log-linear form:

$$lnFM_t = \beta_0 + \beta_1 t + \beta_2 SB + \beta_3 ln\left(\frac{P_t}{CPI_t}\right) + \beta_4 lnRGDP_t + \beta_5 lnNEER_t + \varepsilon_t, \tag{14}$$

where $NEER_t$ is the NEER at Time t. Consistent with the demand theory (Varian, 2010), the sign on β_3 is expected to be negative. In contrast, β_4 is expected to be positive, because rising incomes create a positive influence on food demand (Chambers and Pope, 1992). The sign on β_5 is expected to be negative because an appreciation of domestic currency makes it cheaper to import foreign goods, food included. Although the ARDL does not require *a priori* knowledge of the nature of cointegration among the variables of interest; it does not apply to variables with I (2) and above (Pesaran and Shin, 1999; Pesaran *et al.*, 2001). Unit root tests were employed to rule out this possibility.

3.4 Unit root tests

Three sets of unit root tests – with and without structural breaks on individual time series were used to test for stationarity and to determine their order of integration. The ADFmax unit root test was used to test the null hypothesis that the variable has a unit root against the alternative that a stationary process generated it. The test helped to determine the five variables' order of integration. Prior to conducting the unit root tests, we used the VARSOC command in STATA 12.1 (StataCorp, 2011) to determine the optimal lag order for each of the variables using different information criteria that included the Akaike information criteria (AIC) and the Bayesian information criteria (BIC).

In terms of all food categories, the ADFmax unit root test (Table 1) indicated that three of the series ($lnFM_t$, lnprice and $lnNEER_t$) had a unit root and integrated of Order 1 (i.e. I(1)). Although this test is powerful, its inability to account for structural change against a background of possible structural change in Qatari macroeconomic data makes the ADFmax insufficient on its own. Perron (1989) and Pena (1990) have argued that classic unit root tests are less sensitive in the presence of atypical observations in time series data. Accordingly, the study tested for unit roots in the presence of single and double structural breaks in each of the variables using Zivot and Andrews (1992) (see Table 2) and Clemente *et al.* (1998) unit root tests (see Table 3), respectively, while also estimating optimal structural breakpoints. In keeping with Model C (Equation 5), the null hypothesis for the Zivot–Andrews unit root test was that the series has a unit root, allowing for a single structural break in both intercepts and trend, using a trimming region of 0.15 and 0.85 (see Narayan and Narayan, 2005 for more details). The lack of agreement between the three unit root tests provided further impetus for the need to use an ARDL framework to model the long-run relationship between the variables.

Food group	Variable	t value	Conclusion	Integration order
All Foods	Lnweight_all	−0.332	Unit root	I (1)
	Lnprice_all	−1.312	Unit root	I (1)
	Lnrgdp	−2.810**	Stationary	I (0)
	Lnneer	−1.137	Unit root	I (1)
Meat				
	Lnweight_mt	−3.081***	Stationary	I (0)
	Lnprice_mt	−5.482***	Stationary	I (0)
	Lnrgdp	−2.810**	Stationary	I (0)
	Lnneer	−1.137	Unit root	I (1)
Cereals				
	Lnweight_cer	−2.514**	Stationary	I (0)
	Lnprice_cer	−5.355***	Stationary	I (0)
	Lnrgdp	−2.810**	Stationary	I (0)
	Lnneer	−1.137	Unit root	I (1)
Dairy				
	Lnweight_dai	−1.721	Unit root	I (1)
	Lnprice_dai	−4.446***	Stationary	I (0)
	Lnrgdp	−2.810**	Stationary	I (0)
	Lnneer	−1.137	Unit root	I (1)

Note(s): *p < 0.1; **p < 0.05; ***p < 0.01

Table 1.
ADFmax unit root test

Food group	Variable	Break point	t value†	Conclusion	Integration order
All Foods	Lnweight	2004q1	−61.514	Stationary	I (0)
	Lnprice	2004q1	−5.025	Unit root	I (1)
	Lnrgdp	2007q1	−4.451	Unit root	I (1)
	Lnneer	2002q3	−3.350	Unit root	I (1)
Meat					
	Lnweight	2004q1	−6.241	Stationary	I (0)
	Lnprice	2005q4	−5.930	Stationary	I (0)
	Lnrgdp	2007q1	−4.451	Unit root	I (1)
	Lnneer	2002q3	−3.350	Unit root	I (1)
Cereals					
	Lnweight	2016q1	−4.479	Unit root	I (1)
	Lnprice	2014q2	−7.552	Stationary	I (0)
	Lnrgdp	2007q1	−4.451	Unit root	I (1)
	Lnneer	2002q3	−3.350	Unit root	I (1)
Dairy					
	Lnweight	2004q1	−5.876	Stationary	I (0)
	Lnprice	2016q1	−5.283	Stationary	I (0)
	Lnrgdp	2007q1	−4.451	Unit root	I (1)
	Lnneer	2002q3	−3.350	Unit root	I (1)

Note(s): †5% significance critical value = −5.08

Table 2.
Results of the Zivot–Andrews unit root test with single structural break

Table 3. Results of Clemente, Montañés and Reyes unit root test with double structural break

		Additive outlier model (CLEMAO2)				Innovational outlier model (CLEMIO2)			
Food group	Variable	BP 1	BP 2	*t* value	Integration order	BP 1	BP 2	*t* value	Integration order
All Foods	Lnweight	2003q2	2004q2	−0.608	I (1)	2003q3	2006q2	−71.079	I (0)
	Lnprice	2003q2	2016q4	−0.568	I (1)	2003q3	2016q3	−46.811	I (0)
	Lnrgdp	2003q2	2006q2	−3.435	I (1)	2003q3	2006q3	−6.382	I (0)
	Lnneer	2004q2	2015q2	−5.253	I (1)	2003q2	2014q2	−5.133	I (1)
Meat									
	Lnweight	2004q3	2015q2	−3.916	I (1)	2004q4	2015q3	−8.865	I (0)
	Lnprice	2003q3	2004q3	−3.866	I (1)	2003q3	2004q4	−14.493	I (0)
	Lnrgdp	2003q2	2006q2	−3.435	I (1)	2003q3	2006q3	−6.382	I (0)
	Lnneer	2004q2	2015q2	−5.253	I (1)	2003q2	2014q2	−5.133	I (1)
Cereal									
	Lnweight	2011q1	2015q2	−2.188	I (1)	2003q3	2015q3	−8.489	I (0)
	Lnprice	2007q4	2014q1	−4.108	I (1)	2008q1	2014q2	−5.945	I (0)
	Lnrgdp	2003q2	2006q2	−3.435	I (1)	2003q3	2006q3	−6.382	I (0)
	Lnneer	2004q2	2015q2	−5.253	I (1)	2003q2	2014q2	−5.133	I (1)
Dairy									
	Lnweight	2003q2	2015q2	−3.109	I (1)	2003q3	2015q3	−15.148	I (0)
	Lnprice	2008q1	2015q2	−2.089	I (1)	2008q2	2015q3	−5.517	I (0)
	Lnrgdp	2003q2	2006q2	−3.435	I (1)	2003q3	2006q3	−6.382	I (0)
	Lnneer	2004q2	2015q2	−5.253	I (1)	2003q2	2014q2	−5.133	I (1)

Note(s): BP = break point; 5% significance critical value = −5.49

In terms of aggregate food imports, the results displayed in Table 2 show that while *lnprice, lnneer* and *lnrgdp* are nonstationary, *lnweight* is stationary. In terms of meat and dairy imports, *lnweight* and *lnprice* were found to be stationary. In contrast, *lnrgdp* and *lnneer* exhibited a unit root. In terms of cereal imports, only *lnprice* was stationary.

The Clemente, Montañés and Reyes unit root test (results provided in Table 3) employed both additive outlier (AO) and innovational outlier (IO) models to detect the presence of aberrant observations potentially arising from, respectively, unobservable exogenous and endogenous changes that affect the time series (Tsay, 1986; Pena, 1990). Both models tested the null hypothesis of a unit root (i.e. $H_0 : \rho - 1 = 0$) allowing for a double structural break against the alternative that it is stationary (i.e. $H_0 : \rho - 1 \neq 0$) (See Equation 5).

In the AO model, all the variables were found to be nonstationary. The IO model results were similar to those of the Zivot–Andrews and the Leybourne unit root tests. Overall, however, the Clemente, Montañés and Reyes unit root test results suggest that one would grossly misclassify the order of integration if they relied entirely on traditional unit root tests that ignore structural breaks in the time series data.

Determining the "optimal" breakpoint(s) is crucial given the wide array of feasible possibilities. Accordingly, 27 different models (16 with single and 11 with double structural breaks), were fitted to the data to assess the one with the best fit in the presence of structural breaks. Of the 16 models with a single breakpoint, only six produced promising estimates regarding the sign, magnitude and statistical significance of the ECT and other variables in the ARDL. Only two of the 11 models with a double structural break were suitable for further analysis. The AIC and BIC methods were used to assess all eight candidate models and select the best one.

3.5 Bounds test of cointegration

We used the bounds test of Pesaran *et al.* (2001) to assess the model identified in Section 3.4 and to determine whether there was a statistically significant long-run relationship in levels among the underlying variables. One advantage of the bounds test is that it is applicable irrespective of whether the underlying variables are purely I (0), purely I (1) or mutually cointegrated. Different unit root tests undertaken provided contradicting results; hence, the order of cointegration among the four variables was deemed uncertain.

The bounds test of Pesaran *et al.* (2001) is based on the F statistic whose asymptotic distribution is nonstandard under the null hypothesis of no cointegration among the underlying variables, irrespective of their order of cointegration. There are two sets of asymptotic critical value bounds for regressors, either purely I (0) or purely I (1) under five different deterministic model specifications (See Pesaran *et al.*, 2001 for further details). This study fell under the case of unrestricted intercept and no time trend. The results from the test revealed that for each category of food imports (all foods, meat, cereal and dairy), the null hypothesis of no level relationship between levels of the five variables could not be sustained. In short, there was a statistically significant cointegrating relationship among the four variables. The critical values used were obtained from Kripfganz and Schneider (2018). The results are presented in Table 4.

4. Results and discussion

4.1 Descriptive statistics

Table 5 presents the summary statistics of the variables used in the models. The average quantity of cereals imported into Qatar over the study period was approximately 50 million kg. In addition, during the study period, the country imported an average of 11 million kg of dairy products and 12 million kg of meat. Over the study period, the total food imports

Food group	Test statistic	Significance	Critical values	
All foods	*F* statistic		I (0)	I (1)
	8.513	10	3.482	4.564
		5	4.081	5.267
		1	5.410	6.804
	t statistic			
	−5.552	10	−3.086	−3.798
		5	−3.398	−4.143
		1	−4.013	−4.808
Meat	*F* statistic			
	15.490	10	2.931	3.884
		5	3.414	4.463
		1	4.492	5.740
	t statistic			
	−8.439	10	−3.035	−3.741
		5	−3.354	−4.096
		1	−3.984	−4.781
Cereals	*F* statistic			
	15.220	10	2.948	3.878
		5	3.431	4.453
		1	4.508	5.719
	t statistic			
	−6.309	10	−3.050	−3.757
		5	−3.367	−4.109
		1	−3.993	−4.789
Dairy	*F* statistic			
	27.652	10	3.424	4.638
		5	4.044	5.398
		1	5.444	7.100
	t statistic			
	−9.498	10	−3.022	−3.740
		5	−3.352	−4.012
		1	−4.012	−4.834

Source(s): Critical values from Kripfganz and Schneider (2018)

Table 4.
ARDL Bounds test

Variable	Food category			
	Cereals	Dairy	Meat	All
Quantity (,000)	50899.480 (70286.300) [4.027, 269,000]	11448.810 (15338.800) [8.085, 55,000]	12843.780 (17741.360) [0, 57,900]	145270.800 (255024.900) [0, 707,000]
Price	2.804 (7.266576) [0.532, 61.072]	13.034 (7.699) [1, 39.105]	10.340 (4.186) [2.318, 23.617]	1.162 (2.323) [0, 6.364]
Value (,000)	73325.840 (106,458) [10.951, 369,000]	91164.300 (129423.2) [140.696, 497,000]	124217.900 (189498.1) [0, 552,000]	699270.800 (1,221,583) [0, 3,230,000]
NEER				113.25 (8.576) [58.827, 123.683]
RGDP				0.399 (0.519) [0, 1.589]

Note(s): Standard deviation in parentheses; min. and max. values are in square brackets

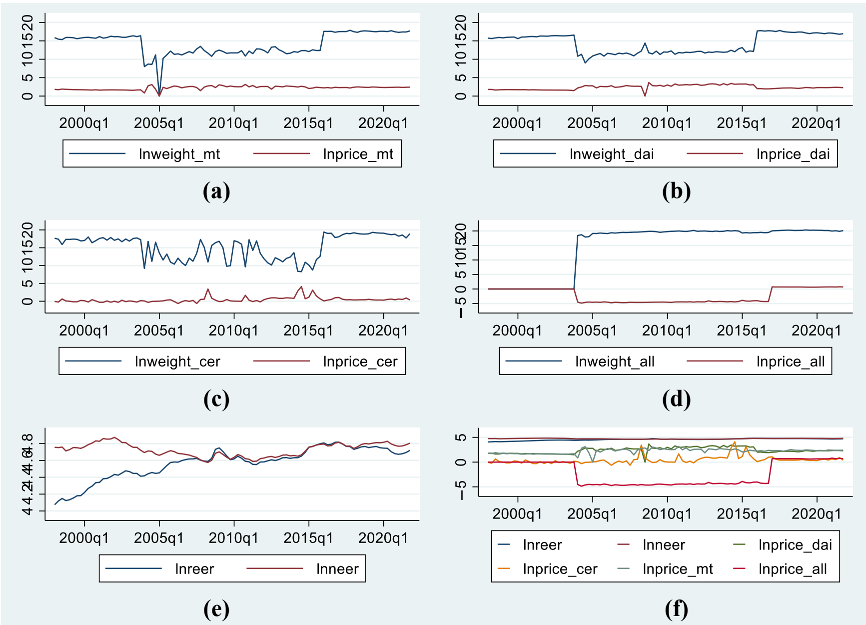
Table 5.
Descriptive statistics

averaged 145 million kg, at an average value of approximately 700 million Qatari Riyals. The nominal and real effective exchange rates averaged 113.25 and 98.04 Qatari Riyals, respectively, to the US dollar. The quarterly per capita GDP growth rate averaged approximately 0.4%.

Figure 1 provide the time series plots of the variables. Plots (a), (b), (c) and (d) show the logs of quantity and price series for meat, dairy, cereals, and all foods, respectively. Plot (e) displays the real and NEERs and (f) describes the relationship between the exchange rates and import prices of the various food categories. In the meat, dairy and cereal plots, an upward shift in the import quantities can be observed, with prices falling slightly after the first quarter of 2015. However, in the all food plot, there is a clear upward shift in the price series and a slight upward shift in import quantity for the same period. This finding indicates that Qatar faced some import difficulties, perhaps resulting from the 2017 embargo.

4.2 ARDL model results

Table 6 presents the estimates from the ARDL model. The study included four models representing Qatar's food imports: all foods (Model A), meats (Model B), cereals (Model C), and dairy (Model D). The models were estimated independently. The sign and significance of the ECT (ADJ.) provide further evidence of the existence of a stable long-run relationship between Qatar's food imports and the set of regressors. The magnitude of the ECT coefficient suggests that Qatar's aggregate food import system corrects itself from the previous period's disequilibrium at a speed of 97% quarterly. In terms of meat, cereals and dairy imports, the system corrects the previous periods' disequilibrium at the speeds of 98%,



Source(s): Qatar Planning and Statistics Authority, (2022); International Monetary Fund, (2022)

Figure 1.
Time series' plots of variables

Variable	Model A	Model B	Model C	Model D
ADJ.	−0.970***	−0.980***	−0.575***	−1.345***
	(0.008)	(0.099)	(0.096)	(0.142)
LR				
$lnprice\_all_{t-1}$	0.071***	1.216	−2.993***	−2.757***
	(0.025)	(0.760)	(0.700)	(0.190)
$lnrgdp_{t-1}$	0.096**	−0.641*	−3.146***	0.226*
	(0.045)	(0.352)	(0.900)	(0.118)
$lnneer_{t-1}$	−5.191***	18.033***	17.865***	1.921
	(0.442)	(4.303)	(6.306)	(1.300)
SR				
$\Delta lnweight_{t-1}$	−0.015*	0.270***	−0.198**	0.472***
	(0.008)	(0.100)	(0.090)	(0.108)
$\Delta lnweight_{t-2}$	−0.063***	0.254***	–	0.490***
	(0.008)	(0.086)	–	(0.095)
$\Delta lnweight_{t-3}$	–		–	0.216***
	–		–	(0.076)
$\Delta lnprice$	0.069***	0.160	−1.721***	−1.191***
	(0.025)	(0.430)	(0.315)	(0.157)
$\Delta lnprice_{t-1}$	–	−1.032*	–	1.884***
	–	(0.515)	–	(0.425)
$\Delta lnprice_{t-2}$	–	−0.686*	–	1.361***
	–	(0.364)	–	(0.330)
$\Delta lnprice_{t-3}$	–	–	–	0.406**
	–	–	–	(0.194)
$\Delta lnrgdp$	0.094**	−0.629*	−1.754**	0.303*
	(0.044)	(0.350)	(0.673)	(0.165)
$\Delta lnrgdp_{t-1}$	–	–	2.044***	–
	–	–	(0.708)	–
$\Delta lnneer$	−2.120**	3.756	11.266	−3.258
	(0.935)	(8.026)	(10.515)	(3.609)
$\Delta lnneer_{t-1}$	2.415**	−10.013	−17.423	−4.737
	(0.963)	(8.997)	(11.291)	(3.695)
$\Delta lnneer_{t-2}$	1.356	−15.281*	−5.006	−9.385**
	(0.944)	(8.231)	(11.030)	(3.537)
$\Delta lnneer_{t-3}$	–	−13.558*	−29.768***	–
	–	(7.436)	(10.269)	–
Covid	−0.128	0.072	−0.099	−1.202***
	(0.078)	(0.566)	(0.887)	(0.308)
Blockade	0.288***	1.583***	0.641	0.362
	(0.102)	(0.536)	(0.869)	(0.281)
Datevar	0.016***	0.043**	0.026*	0.131***
	(0.002)	(0.019)	(0.014)	(0.017)
stbrkall2004q4	−1.165***	–	–	–
	(0.143)	–	–	–
sbmt2005q1	–	−11.842***	–	–
	–	(1.578)	–	–
sbcer2014q2	–	–	−1.905	–
	–	–	(1.918)	–
sbdai2016q1	–	–	–	2.978***
	–	–	–	(0.574)
_cons	39.480***	−81.660***	−43.662**	−11.993
	(2.012)	(19.903)	(18.642)	(7.827)
Adj- R^2	1.00	0.78	0.55	0.83
AIC optimal lags sel	(3, 0, 0, 3)	(3, 0, 0, 3)	(2, 0, 2, 4)	(4,4,0,3)
N	72	72	72	72

Note(s): $*p < 0.1$; $**p < 0.05$; $***p < 0.01$; standard errors are in parentheses

Table 6.
Estimates from the ARDL model

57.5% and 103.5%, respectively. Such a high speed of adjustment unequivocally reflects how sensitive and, therefore, vulnerable, Qatar is to potential short-run distortions in food import supply.

In line with the *Le Chatelier's* principle, the long-run elasticities are larger than their short-run counterparts. This finding indicates that the model is theoretically consistent (Fernandez-Cornejo, 1992). The long-run price elasticity of import demand is highly significant and greater than unity in two of the four models: Models C and D (i.e. cereals and dairy). In Models A and B (the all foods model and the meat imports model) the price elasticity is positive and less than unity in the long run. The greater than unity elasticity of the two models indicates the high sensitivity of Qatar's cereals and dairy imports to relative price changes. In the short run, only Models C and D exhibited a greater than unity price elasticity. As expected, in practice, this sensitivity was lower in the short run in absolute terms (1.721) compared with the long run (2.993) for cereal imports, and in Model D where the figures were 2.757 and 1.191, respectively. This phenomenon is intuitive, as consumers find it more difficult to switch their tastes and preferences in the short run compared to the long run, either because of a lack of better alternatives or out of sheer habit inertia (Mankiw, 2004). All else being equal, a 1% increase in import price would reduce the quantity of cereal imports by 1.721 and 2.993% in the short and long run, respectively. In terms of dairy imports, a 1% increase in import price would lead to 1.191 and 2.757% decrease in imports in the short and long run, respectively. In their studies, Agbola and Damoense (2005) and Nguyen and Jolly (2013) reported long-run price elasticities of import demands of -1.7 for chickpeas in India and 1.5 for seafood in The Caribbean, respectively, findings of which tally with our results.

The income elasticity of import demand for all foods during the study period was positive and less than unity. As expected, this result indicates that food is a necessity in Qatar. The income elasticity was slightly higher in the long run (0.096) than in the short run (0.094). Hence, a 1% increase in real GDP would lead to a 0.096 and 0.094% increase in demand for all food imports, *ceteris paribus*. This finding suggests that an increase in short-run food import prices would most likely affect the lifestyle of Qatar's residents in the long run. A similar result was observed in the dairy imports model where the income elasticity was positive. However, in the meat and cereals import models, the coefficient of GDP was negative, in both the short and the long run, indicating negative income elasticity. This finding indicates that as their incomes rise, Qataris replace meat and cereal with other sources of protein and carbohydrates, thereby leading to a fall in the demand for the former.

In the short run, the coefficient on NEER was not statistically significant and did not have the expected negative sign in the meat, cereals and dairy import models. This finding could be attributed to Qatar's fixed exchange rate policy that pegs Qatari Riyal to the US dollar at 3.64:1. As observed in this study, the effective exchange rate may be insensitive to short-term domestic fluctuations thereby having no effect on food imports. In the long run, however, the NEER changes in tandem with the global strength of the US dollar that translates into a huge negative effect on Qatar's aggregate food imports. Accordingly, a 1% increase in the NEER would reduce Qatar's aggregate food imports by 5.191%, meat imports by 18.033% and cereals imports by 17.865%.

The results also provide evidence that the 2017 blockade led to a 0.288% increase in aggregate food imports. In terms of meat imports, there was a 1.583% increase in its imports. However, no statistically significant effect was registered on Qatar's cereal and dairy imports. The study also modeled the impact of the COVID-19 pandemic on food imports: the ongoing pandemic has generated global supply disruptions revealing Qatar's vulnerability to crises (Hassen *et al.*, 2022). The pandemic significantly reduced dairy imports by 1.202%. Although statistically insignificant, a similar effect was found for cereal imports and aggregate food imports. The results also revealed that structural breaks have significant effects on food imports in all, but the cereal imports model (Model C).

4.3 The Toda–Yamamoto causality test

Having established the existence of a long-run relationship between aggregate food imports and various regressors, it was essential to ascertain the causal relationship among them. According to Granger (1969), a causal relationship in the short run exists between two stationary variables if one variable can be better predicted by its own lags and those of the other variable. If the coefficients on past values of the predicting variable are statistically significant, then that variable is said to Granger-cause the predicted variable and vice versa. The only requirement for Granger causality to exist is that any pair of variables tested should be cointegrated (Granger, 1988). The mixture of I (0) and I (1) variables necessitated the use of the Toda and Yamamoto (1995) causality test. These authors argue that undertaking a regular Granger causality test on variables where there is uncertainty about their order of integration, cointegration and/or trend stationarity violates the standard asymptotic theory due to the presence of nuisance parameters. The Toda and Yamamoto (1995) test is essentially a Granger causality test augmented with extra lags on each variable that represent the maximal integration order. These lags are explicitly captured in the set of exogenous regressors in the VAR to fix the asymptotics. The null hypothesis tested was that a particular variable, X_t, does not Granger-cause another variable, Y_t, against the alternative that at least one coefficient on the lagged values of X_t is not equal to zero.

The Toda–Yamamoto causality test results for each of the four food categories considered in this study (i.e. all foods, meat, cereals and dairy) are presented in Table A1 in the online Appendix. Based on the results, the null hypothesis that aggregate food imports in Qatar are not Granger-caused by the other three variables was not sustained. In particular, the study found that the NEER had a statistically significant causality on Qatar's meat, cereal and dairy imports and that this causality was unidirectional. This finding emphasizes the importance of a NEER on Qatar's food imports, particularly given that the country's food security is heavily dependent on imports. The apparent statistical significance of the result for real income can be plausibly attributed to temporal aggregation than to actual instantaneous causality. Additionally, there may exist an unobserved missing variable that causes the apparent significance of Granger causality among the control variables in the real income equation, a point Granger notes (1988, p. 208). In contrast, there is no statistically significant Granger causality for the real unit import price (for meat and dairy imports) and the NEER. In short, Qatar's meat and dairy imports are mainly driven by the effective exchange rate.

4.4 Impulse response functions

Figures 2–5 present the response surfaces, showing the response of Qatar's meat, cereals, dairy and the aggregate of all food imports to exogenous shocks over five periods. As shown in Figure 5, food imports would decrease marginally if the unit price were shocked by one standard deviation in the first period. However, the study found this decrease was not statistically significant. A shock of one standard deviation on the NEER would lead to a statistically significant decrease (2.5%) in food imports during the first period. The fall in imports would continue through to the third period. While a similar shock on real income would marginally raise the food imports in the first and second periods, the study found that the change was not statistically significant. In other words, only the NEER has a significant impact on Qatar's food imports. In terms of meats, cereal and dairy imports, however, a shock of one standard deviation in the NEER would result in an upward shock in imports, except for cereal imports where there was an initial downward shock in the first period (Figure 3).

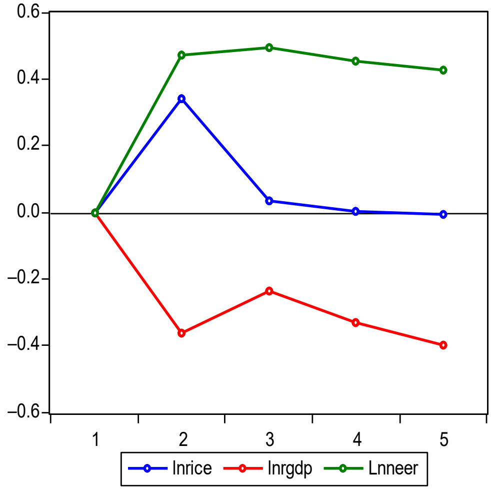


Figure 2.
Meat impulse response

Source(s): Qatar Planning and Statistics Authority, (2022); International Monetary Fund, (2022)

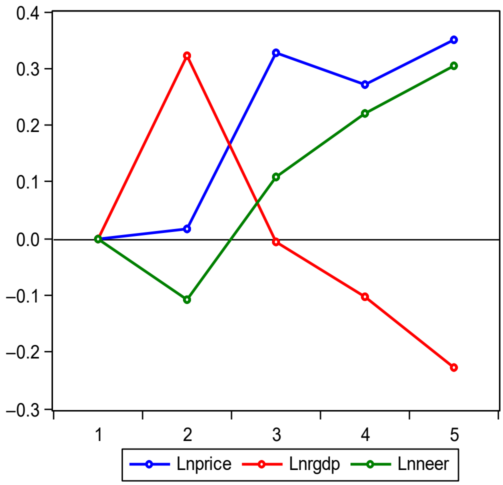
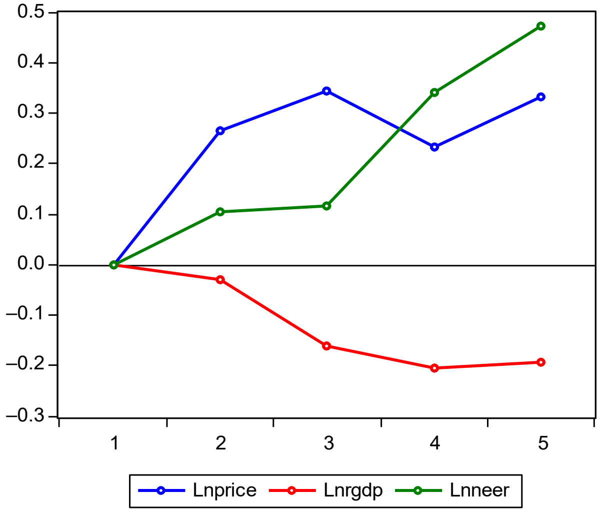


Figure 3.
Cereal impulse response

Source(s): Qatar Planning and Statistics Authority, (2022); International Monetary Fund, (2022)

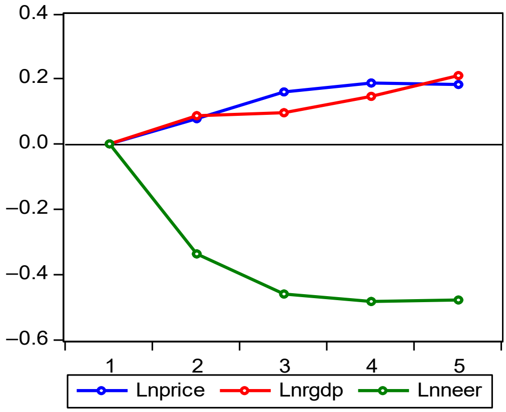
5. Conclusion

This study employed Pesaran *et al.* 's (2001) bounds testing procedure within an ARDL framework to examine the long-run relationship between aggregate food import demand, import price, real income and the NEER for Qatar. It demonstrates the need to go beyond casual unit root testing, particularly given the insensitivity of classical tests to possible regime change in time series data. The results show a unique equilibrium relationship among the four chosen variables. Further, as predicted by the *Le Chatelier's* principle, the results



Source(s): Qatar Planning and Statistics Authority, (2022); International Monetary Fund, (2022)

Figure 4.
Dairy impulse response



Source(s): Qatar Planning and Statistics Authority, (2022); International Monetary Fund, (2022)

Figure 5.
All food impulse response

demonstrate that the long-run elasticities are larger than their short-run counterparts, thereby indicating a theoretically consistent model.

The long-run price elasticity of import demand for all foods (0.071) and for meat imports (1.216) indicates a positive relationship between imports and import tariffs. While this finding is contrary to theory, it reflects Qatar's dependence on food imports that perhaps is made possible by the country's growing per capita income. This result is somewhat consistent with Suanin's (2021) observation that growth in income neutralizes the negative effect of price increases. It is also consistent with Hoang's (2018) work in Vietnam. The author found that at higher incomes demand for food is less responsive to price changes. For cereals and dairy imports, we found that a 1% increase in their respective import tariffs would result in a 2.993 and 2.757% reduction in cereals and dairy imports. Qatar is part of the GCC customs union and is thus subject to a common external tariff that may change over time and affect its food imports. This finding is consistent with Niemi (2018) claim that the geography of the importing country can impact upon the elasticity of import demand for some commodities. Hatab and Surry (2022) observed a similar interaction between the import price and the quantity of potatoes imported from Egypt. The current study also found that in the short and long run, income has a negative, albeit marginal, effect on meat and cereals import demand. The effect was larger on cereal imports relative to meat. This finding is consistent with the Engel's law. According to this law, individuals tend to allocate relatively larger proportions of their additional income to nonfood expenditure, hence spending less on food, resulting in lower demand for food. Bairagi *et al.* (2020) observed a similar relationship between income and the demand for rice in Vietnam. The study found that for the study period the NEER has a long-run negative effect on Qatar's aggregate food imports. However, when analyzed in isolation, the NEER had a positive impact on both the meat and cereal imports.

The short-run dynamics indicated a high adjustment speed, implying that, in the case of an exogenous shock, it would take a relatively short time for food imports to revert to equilibrium. The impulse response functions revealed that such shocks were primarily associated with Qatar's NEER, perhaps precipitated by fluctuations in the dollar-denominated global economy where Qatar plays a major role in hydrocarbon trade. These issues suggest Qatar's high vulnerability to distortions in its food import supply chain.

The Toda and Yamamoto (1995) causality test confirms a unidirectional causal effect on the three regressors on Qatar's food imports, with the NEER being the most important. Accordingly, an exogenous shock of one standard deviation on this variable would reduce Qatar's food imports by approximately 2.5% during the first period. They would recover after the third period. In an economy with a pegged exchange rate policy, externally driven volatility in the effective exchange rate is likely to pile pressure on food imports, thereby jeopardizing the country's food security. As demonstrated by the minimal impacts on food imports during both the 2017–2021 economic embargo and the COVID-19 pandemic, Qatar has deployed its vast fiscal resources to mitigate the effects of adverse external shocks. However, the long-term sustainability of such efforts remained to be seen, especially with the growing global shift away from fossil fuels as an energy source.

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(The Appendix follows overleaf)

Appendix

Equation	Excluded	All foods model χ^2 statistic	Meat χ^2 statistic	Cereals χ^2 statistic	Dairy χ^2 statistic
Lnweight	Lnprice	0.032 (1)	1.013 (1)	0.332 (1)	6.745 (1) ***
	Lnrgdp	0.577 (1)	7.560 (1) ***	3.820 (1) **	14.129 (1) ***
	Lnneer	2.640 (1)	15.454 (1) ***	19.940 (1) ***	34.365 (1) ***
	All	3.305 (3)	21.967 (3) ***	29.692 (3) ***	53.521 (3) ***
Lnprice	Lnweight	8.472 (1) ***	0.282 (1)	1.974 (1)	1.531 (1)
	Lnrgdp	4.524 (1) **	2.183 (1)	5.174 (1) **	4.771 (1) **
	Lnneer	6.118 (1) **	15.084 (1) ***	7.910 (1) ***	36.693 (1) ***
	All	26.529 (3) ***	34.221 (3) ***	15.211 (3) ***	49.593 (3) ***
Lnrgdp	Lnweight	13.180 (1) ***	5.499 (1) **	0.369 (1)	7.320 (1) ***
	Lnprice	12.792 (1) ***	1.477 (1)	1.652 (1)	6.890 (1) ***
	Lnneer	0.120 (1)	1.138 (1)	0.120 (1)	1.591 (1)
	All	14.716 (3) ***	9.067 (3) **	1.742 (3)	8.767 (3) **
Lnneer	Lnweight	0.188 (1)	0.350 (1)	0.955 (1)	1.046 (1)
	Lnprice	0.228 (1)	1.482 (1)	13.706 (1) ***	1.709 (1)
	Lnrgdp	16.220 (1) ***	24.139 (1) ***	9.265 (1) ***	16.004 (1) ***
	All	17.758 (3) ***	30.401 (3) ***	40.93 (3) ***	26.528 (3) ***

Note(s): *p < 0.1; **p < 0.05; ***p < 0.01; degrees of freedom are in parentheses

Table A1.
Results of Toda–Yamamoto causality test

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