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Aggregate Import Demand Function For Saudi Arabia: An Error Correction Approach

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Abstract

The paper attempts to analyze Saudi Arabia's aggregate demand for imports. Cointegration analysis and two different approaches of Error Correction (Engle-Granger's and Hendry's error correction models) were applied to contrast their performance in fitting the Saudi data. It is found that in both models, domestic price, import price, and income are important in determining the import demand. The empirical results show that aggregate import demand tends to be income and price elastic. Moreover, the results give an evidence that Engle-Granger approach outperforms the other model in terms of having the smallest ex-post forecast errors.

Key words: *Demand; Import; ECM; Saudi Arabia*

Introduction

In studies of international trade, import demand functions are often estimated using standard econometric procedures. Among several important econometric issues that have been discussed lately, is the stability of trade relationships over time. Nelson and Plosser (1982) provided a widely known empirical testing for the stability of economic time series data. They concluded that most economic time series are not stationary. Moreover, another important study by Goldstein and Khan (1985) has suggested that trade relationships are subject to either gradual or sudden changes over time. They argued that the gradual changes are due to the process of economic development or as the result of changes in government trade policies, whereas sudden changes are due to fluctuations in exchange rates, or large oil price increases that alter the basic demand or supply relationship (p. 1073).

Recent developments in econometrics literature have provided a method of integrating the long-run relationships between economic variables with short-run dynamic adjustment processes for non-stationary variables. Such developments include the contribution of Engle and Granger (1987), Johansen and Juselius (1992) and Johansen (1988, 1991, 1996). Their efforts in testing for cointegration have provided some methods of integrating the long-run relationships between economic variables with short-run dynamic adjustment processes. This progress has revived the interest of many scholars to search for a long-run relationship for many economic variables. In this paper, these techniques are applied to derive a long run equilibrium demand function of imports and error correction model using Johansen procedures.

This study is intended to apply cointegration and two error correction techniques to estimate short and long run elasticities in import demand for Saudi Arabia. The paper is set up as follows; the model and its theoretical underpinnings are laid out in section 2, section 3 provides a fairly detailed discussion of the econometric procedures used in the analysis of cointegration, section 4 presents the empirical results. The paper concludes in section 5 with a summary and assessment of the findings.

Modeling Import Demand

The standard specification of the import demand model is similar to any other demand model. It treats the import quantity demand as dependent variable and import price and income as independent variables. The basic function of the import demand model, which is referred to as the imperfect substitute model of trade (Goldstein and Khan, 1985), is of the following form:

$$M_t^d = f(Y_t, PD_t, PM_t)\,, \tag{1}$$

where M_t^d is the quantity of total imports demanded in time period (t), Y_t is the Gross Domestic Product (GDP) in period (t), PD_t is the price of domestic goods in period (t), and PM_t is the price of imports in period (t). This general form function indicates that demand of imports can be explained by the income, domestic prices and imports prices. It is assumed in this function that imports are normal goods that are substitutable for domestic goods. In

this specification, the equilibrium price and quantity are determined by the interaction of supply and demand forces. Murray and Ginman (1976) argued

that models that have this kind of specification involve an identification problem. This problem is usually solved in international trade studies by assuming that supply elasticity is infinite. This crucial assumption simplifies the model and reduces it statistically to a single equation.

Marquez and McNeilly (1988) raised some key issues in modelling import demand such as the selection of explanatory variables, the choice of functional form, and the characterization of dynamic adjustments. The most widely known specification of import demand equations that considers these issues is of the following log linear form:

$$\ln M_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln PM_t + \beta_3 \ln PD_t + u_t, \qquad (2)$$

where

M_t = Total imports of Saudi Arabia in time period (t),
Y_t = GDP of Saudi Arabia in time period (t),
PM_t = Unit Value (price) of imports of Saudi Arabia in time period (t),
PD_t = Price index of domestically produced goods of Saudi Arabia in time period (t),
u_t = Error term.

In the above specification, we have used two separate price terms instead of the relative price ratio, to capture the price effects on imports. Murray and Ginman (1976) argued that the relative price specification of the traditional import demand model is inappropriate for estimating aggregate import demand parameters. They suggested a simple modification of the traditional import demand equation that estimates the impact of import and import competing prices separately. Moreover, Urbain (1993) argued that the use of two separate price terms is preferable to the use of one term. He offered some economic and econometric reasons that support this argument.

In equation (2), the price terms enter independently; i.e., we do not restrict the coefficients of domestic and import prices to have the same magnitude, but opposite signs, so that ($-\beta_2=\beta_3$). The imposition of such a restriction means that the imports demand equation is homogeneous of degree zero in prices. This allows researchers to estimate the model with relative prices instead of price levels. However, if this type of restriction is not valid, it can lead to bias in estimated coefficients as shown by Hynes and

Stone (1983). They argued that the expected value of invalid restricted coefficient could very well lie outside the range of unrestricted coefficients

and might reverse signs. Urbain (1993) has extended the argument against the use of relative prices. He stated that modeling the dynamics of imports demand using relative prices implies identical dynamic response of imports

to changes in imports prices and domestic prices. This situation is difficult to qualify, as economic agents use different information sets to form their expectations about domestic and foreign (imports) prices. Also, the latter price embodies some expectations about the exchange rate.

Most of empirical trade work has been confined to the estimation of functions for imports (Goldstein and Khan, 1985, pp 1076). The basic specification of import demand, given by equation (2), is very popular in international trade studies. In most cases, the formulation of import demand equation is given in terms of logarithms. The logarithms formulation is preferable in modeling import demand functions for two reasons. First, it gives direct estimation of import elasticity. Second, it allows imports to react proportionally to a rise and fall in the explanatory variables (Khan 1975).

Assery and Peel (1990) estimated a simple aggregate import demand function for five developed countries (Canada, Japan, United Kingdom, United States and West Germany) using quarterly data for the period (1972-1986), using two different estimation procedures. Their study concluded that previous estimates of long-run income and price elasticities of imports were subject to substantial upward bias because those estimates were based on standard econometric procedures.

Stone (1979) presented the first comprehensive study to estimate price elasticities of demand for the imports and exports of the United States and some other developed countries. Warner and Kreinin (1983) used the above standard specification to estimate import demand functions for 19 countries and two time periods: 1957-1970 and 1972-1980. In addition to the conventional income and price variables, their study estimated an unrestricted lag structure of the effect of price and exchange rate variation on imports. Their estimates have been corrected for serial correlation using the technique of Cochrane-Orcutt. One of the main shortcomings of the above standard modeling of imports is that it does not account for cross-price elasticities. They argued that these elasticities are not considered, because of the difficulty in estimating them statistically. Therefore, they developed and

applied a model that calculates cross elasticities of U.S. demand with respect to the prices of imports from different countries. Finally, Goldstein and Khan (1985) summarized the literature on import price and income elasticities that

have emerged in previous empirical studies, and concluded that import price elasticities for the short-run demand are considerably smaller than those for the long run.

Econometric Procedures

The basic assumption underlying the standard estimation procedures is that the time series are stationary, in the sense that the mean and variance are independent of time. However, many economic time-series are not stationary and change over time (Nelson and Plosser, 1982). This means that, as time goes on, the mean and variance tend to move away from any given values. Non-stationarity is usually removed by taking first differences (Box and Jenkins, 1970). However, this also results in removing out the long-run characteristics of the data, thereby making the model capable of explaining only short-run effects.

Although many time series may tend to trend up or down over time in a non-stationary behavior, a group of them might drift together. If there is a tendency for some variables to hold a linear relationship over long periods of time, then cointegration analysis can be used to find out this long-run equilibrium relationship.

Testing for Non-Stationarity and Order of Integration

A time series is stationary if its mean, variance, and autocovariances are independent of time. Assume that M_t is a time series that is generated by a process that follows a first order autoregressive model:

$$M_t = \rho M_{t-1} + \varepsilon_t , \tag{3}$$

where ε_t is a white noise that represents a sequence of independent error terms. To test for non-stationarity of the M_t series, we test the null hypothesis of a unit root that is (ρ=1) versus the alternative hypothesis of ($|\rho| < 1$)[(1)]. Rejecting the null hypothesis means the series M_t is stationary and integrated of order zero. The most commonly used test in the literature is

[(1)] In most economic series, it is conventional to assume that ($\rho \leq 1$), for more details see (Perman, 1991).

the Dickey Fuller (DF) test. Dickey and Fuller (1979) showed that the well-known (t) and (F) statistics are inappropriate for testing the null hypothesis of a unit root. They constructed the DF tables that have to be employed when investigating the existence of unit roots. Later, and as a result of some major criticisms, Dickey and Fuller (1981) presented another powerful test known as Augmented Dickey- Fuller test (ADF). This test accounts for the fact that the residuals are estimates of the true disturbances, and the possibility of having autocorrelation in the error process term. The ADF test is presented by adding lagged independent variable values when the disturbance term (ε_t) is not a white noise[(2)].

The next step is to find the order of integration of the time series M_t. In other words, to find the number of times the series need to be differenced to achieve stationarity. Most economic series are known to be integrated of order one. Therefore, it is usually appropriate to start testing the hypothesis that the order of integration is one, that is $\Delta M_t \sim I(0)$. In this case the new DF regression equation is written as

$$\Delta M_t = \delta \ \Delta M_{t-1} + \varepsilon_t \,. \tag{4}$$

Again, the rejection of the null hypothesis δ=0, implies that the time series ΔM_t is stationary and integrated of degree one, whereas the acceptance of the null hypothesis implies that the time series ΔM_t is non-stationary, and could be integrated of order higher than one. Theoretically, this process should continue until an order of integration is established.

Cointegration Analysis

Cointegration analysis refers to the process of getting equilibrium or long-run relationships among non-stationary variables. The idea is that although the variables are non-stationary, a linear combination of them may be stationary, given that all variables are integrated of the same order (Engel and Granger, 1987). The vector that links the variables in the long-run relationship is called the cointegrating vector.

The most common procedure to test for cointegration is the Engle-Granger two-step estimation technique (EG). The first step in this method

(2) Some other alternative tests for unit roots are explained in Charemza and Deadman (1997).

implies fitting the long-run relationship in levels by OLS and using the resulted residuals to test the hypothesis of cointegration by applying the DF

test. If the hypothesis of cointegration is accepted, then there exists an error correction representation (Engle and Granger, 1987). The second step is to construct the Error Correction Model, which represents the short-run dynamics. Another procedure to test for cointegration is developed by Johansen and Julesius (1988, 1992), and it is known as the maximum

likelihood (ML) approach. This method estimates and tests for multiple cointegrating vectors (multivariate cointegration). It applies the analysis of the vector auto-regressive (VAR) model where all variables are treated as endogenous.

The Error Correction Model (ECM)

The most important finding of cointegration analysis is "*The Granger Representation Theorem*". This theorem states that if a set of variables is cointegrated of order 1,1 [CI (1,1)], there exists a valid error-correction representation of the data. Engle and Granger (1987) provided a principal feature of the cointegrated variables in that their time paths are influenced by the deviation from the long-run relationship, given that cointegration implies error correction representation. That is, a cointegrated system can always be represented by an ECM, which is described as follow:

$$\Delta LM_{it} = \alpha_0 + \sum_{i=1}^{n} \alpha_{1i} \Delta LM_{t-1} + \sum_{i=0}^{n} \alpha_{2i} \Delta LY_{t-1} + \sum_{i=0}^{n} \alpha_{3i} \Delta LPD_{t-1} + \sum_{i=0}^{n} \alpha_{4i} \Delta LPM_{t-1} + \gamma u_{t-1} + \upsilon_i , \quad (5)$$

where the first difference of imports is a function of lagged imports value, current and lagged values of the independent variables, and the lagged value of the long-run disturbance term (u_{t-1}). The parameter (γ) describes the short-run adjustment and indicates the speed of adjustment towards the long-run equilibrium state. An ECM is very appealing, because it includes the short-run and the long-run effects.

To show the above merits of the ECM, we use the following simple equilibrium equation for illustration:

$$M_t = \alpha + \beta X_t ,$$

where M is a dependent variable and X is a vector of independent variables. In case of quilibrium the balance (M-α–βX) will equal zero. However, (M-

$\alpha - \beta X$) will be non-zero when disequilibrium occurs. More precisely, this quantity measures the extent of disequilibrium between M and X and hence

is known as the disequilibrium error. In case of disequilibrium, the typical form is as follow:

$$\Delta M_t = \delta_0 \Delta X_t - \mu(M_{t-1} - \alpha - \beta_1 X_{t-1}) + u_t, \qquad (6)$$

where Δ represents the first difference. Equation (6) shows that the change in M depends on the change in X and the lagged value of the disequilibrium error, which implies that when M_{t-1} is greater than the disequilibrium value, the value of M will be decreased in the subsequent period. The model thus measures how the value of M is corrected for the disequilibrium error and hence is called the ECM. From equation (6), it is clear that δ_0 and b_1 measures the short-run and long run parameters. In practice, higher lag orders are added to equation (6) as explanatory variables to make u a white noise. When higher order lagged variables are introduced, equation (6) is modified into the following form:

$$\Delta M_t = \sum_{i=1}^{k-1} \Psi_i \Delta M_{t-1} + \sum_{i=0}^{k-1} \delta_i \Delta X_{t-i} - \mu(M_{t-k} - \alpha - \beta X_{t-k}) + u_t, \qquad (7)$$

There are two approaches to estimate the ECM: the Engle-Granger two- step procedure and Hendry’s type of testing down. Asymptotically, both methods should yield a valid ECM, and hence arrive at the same estimated model (Granger, 1986). However, in small samples an ECM obtained from Engle-Granger procedure may not necessarily correspond to the type of model derived from Hendry’s testing down. In view of this, one cannot say a priori which method is preferable. As we have a finite sample, both approaches will be employed in this study and their forecasting errors will be compared in order to differentiate their performance in modeling the import demand behavior for Saudi Arabia.

Empirical Results

In this study, the aggregate import data for Saudi Arabia were obtained from Direction of Trade Statistics (DOTS). The data for the other variables, however, were obtained from International Financial Statistics (IFS). These variables include gross domestic product (GDP), price of imports (PM), price of domestically produced goods (PD). The nominal values of all variables have been converted into real ones using the appropriate GDP deflator and price indices. The data were collected annually covering the time period from 1968

to 1998. However, the price of imports, which is usually replaced by the import unit value, was not available for Saudi Arabia. Therefore, since the largest share of Saudi imports comes from industrialized countries, the export unit value for the industrialized countries was used as proxy for that variable.

Stationarity and Unit Root Test

The Augmented Dickey-Fuller (ADF) unit root test was applied to the logarithms of the four time series employed in the study (LM_t , LY_t , LPM_t , and LPD_t) with and without time trend. Table (1) reports the empirical results of the (ADF) unit root tests. The results indicate that non of the reported t-statistics for LM, LY, LPM, and LPD were close to the 5% critical value for the "t-statistic". This means that the null hypothesis of a unit root cannot be rejected for the levels of each variable. Therefore, it is concluded that the LM_t, LY_t, LPM_t, and LPD_t, are non-stationary series. The same results are obtained when the equations include a time trend term. These results present evidence that each time series is integrated of order one; that is each series is I(1). This supports out earlier conjecture that they are non-stationary time series.

Table (1): ADF Tests for Stationarity of the Time Series

Variable	Level		First difference	
	Without trend (-2.9798)	With trend (-3.7522)	Without trend (-2.9850)	With trend (-3.6027)
LM	-2.1627	-2.8360	-3.5679	-3.4634
LPD	-2.9376	-2.9830	-2.9853	-3.6097
LPM	-1.8504	-1.9565	-3.3919	-3.7944
LY	-1.8868	-2.7874	-3.8839	-3.8460

The 5% critical values were obtained from Mackinnon (1991) and are given between parantheses.

Cointegration Analysis

In order to choose the optimal lag length, we tested down from the general 4 lags system. The Schwarz Baysian Criterion (SBC) and the Akaike Information Criterion (AIC) suggested different order of VAR. However, we used one lag as suggested by (SBC) and also because we have annual data.

Table (2) shows the Johansen likelihood ratio statistics for determining the number of cointegrating vectors r, using the maximal eigenvalues test (λ-Max test) and the trace test.

Table (2): Cointegration with Restricted Intercepts and no Trends in VAR (1) LR Test based on Maximal Eigenvalue and LR Test based on Trace of Stochastic Matrix of Series M, PD, PM, and Y

Null	Eigenvalue	Maximal Test	5% Critical Value	Trace Test	5% Critical Test
r=0	0.82979	44.26	28.27	87.02	25.80
r≤1	0.69125	29.38	22.04	42.76	19.86
r≤2	0.26.33	7.53	15.87	13.38	13.81
r≤3	020837	5.84	9.16	5.84	7.53

The source of the critical values is Pesaran & Pesaran (Microfit 4, 1997), see also Pesaran, et al. (1997).

Starting with the null hypothesis of zero cointegrating vector (r = 0), followed by the tests for (r≤1), and (r≤2), the λ-Max test shows that the hypothesis of zero cointegrating vectors is rejected at the 5% level of significance. Thus, the results of the Johansen-Juselius cointegration tests

suggest that there exists one cointegrating vector at the 5% level of significance, which indicates that there exists a long-run stationary relationship between the four variables. The estimated cointegrating vector of demand for Saudi Arabia imports is given by

$$LM_t = -9.48 + 2.60LY_t + 1.79LPD_t - 0.87LPM_t \,. \qquad (8)$$

All three coefficients have correct signs and significant values. The test for a *unit* income elasticity gives a χ^2 (1) = 9.18 [p = 0.010], indicating that the elasticity of income in Saudi Arabia is more than one. The estimated cointegrating vector shows also a long-run domestic price elasticity of 1.79. The test for *unit* domestic price elasticity gives a χ^2 (1) = 31.09 [p < .0001], which indicates that we can reject the hypothesis of a domestic price elasticity of one and conclude that it is more than one. The same estimated vector gives import price elasticity of -0.87. The test for *unit* price elasticity gives a χ^2 (1) = 1.40 [p = 0.496], which means that we cannot reject the hypothesis that the price elasticity is (- 1)[3], and we conclude that it equals one.

[3] However, we tested for the restriction of zero for all three variables and it was rejected for each one, with χ^2(1) = 12.17 [p < 0.0001] for income, χ^2(1) = 8.67 [p = 0.003] for domestic prices, and χ^2(1) = 7.104 [p = 0.008] for import prices.

Estimation of the Error Correction Model
Engle Granger EC Approach

Having obtained the values of the long-run parameters, we can proceed to the second step of the Engle-Granger ECM by feeding those values into the disequilibrium error of Equation (9). The equation to be estimated is:

$$\Delta LM_{it} = a_0 + \sum_{i=1}^{k-1} a_i \Delta LM_{t-i} + \sum_{i=0}^{k-1} b_i \Delta LPD_{t-i} + \sum_{i=0}^{k-1} c_i \Delta LPM_{t-i}$$

$$+ \sum_{i=0}^{k-1} d_i \Delta LY_{t-i} - \gamma [LM_{t-k} - \beta_1 LY_{t-k} - \beta_2 LPD_{t-k} + \beta_3 LPM_{t-k}] + \nu_t . \quad (9)$$

In estimating the above equation, the initial step is to set the value of k and then successively delete the most insignificant differenced variables. We have tried various values of k, with a maximum of 3, while estimating the above equation. The best performing model that we have obtained (k=1) is summarized in Table (3), which shows the estimates of equation (9) for the period 1971-1997. From a statistical point of view the results depicted in this table are satisfactory. The R^2 is reasonably high and there is no sign of residual serial correlation or autoregressive conditional heteroscedasticity in the model. Since the model passed all the diagnostic tests that we have conducted, this indicates that our choice of k is acceptable.

Table (3): Estimates of Engle-Granger's ECM

Variable	Coefficient	Value	t-statistic
Constant	α_0	-0.019	-0.59
ΔLM_{t-1}	a_0	0.210	0.11
ΔLPD_t	b_0	0.495	1.19
ΔLPM_t	c_0	-0.827	-2.01
ΔLY_t	d_0	1.216	2.28
ECM_{t-1}	γ	-0.617	-3.58
$R^2 = 0.79$, F (5,19) = 19.33 [p < 0.0001]			
Normality Test: $\chi 2(2) = 1.475$ [p = 0.478]			
Durbin-Watson statistic = 1.68			
Serial Correlation LM: $\chi 2(1) = 2.335$ [p = 0.126]			
LMARCH: $\chi 2(1) = 0.590$ [p = 0.442]			

Notes: LM and LMARCH are Lagrange multiplier tests for first order autocorrelation and first order autoregressive conditional heteroscedasticity, respectively.

The short–run elasticities of income, consumer price and import price are given by the estimated coefficients on the zero lag terms. These estimates are:

$$\varepsilon_{Y}^{SR} = 1.22, \qquad \varepsilon_{PD}^{SR} = 0.50, \qquad \varepsilon_{PM}^{SR} = -0.83.$$

As expected, all short-run estimates are lower than their corresponding long-run estimates. Also, the estimated value of the adjustment parameter γ [the coefficient of error correction term lagged one period] is –0.617, which means that, if demand is 1% out of equilibrium, a 61.7% adjustment towards equilibrium will take place within the first year.

Hendry's ECM

Instead of applying equation (9) directly, Hendry's ECM normally starts with an unrestricted version of the ECM (Hendry, et al., 1984) as given by the following equation:

$$\Delta LM_{it} = a_0 + \sum_{i=1}^{k-1} a_{i1}\Delta LM_{t-i} + \sum_{i=0}^{k-1} a_{i2}\Delta LPD_{t-i} + \sum_{i=0}^{k-1} a_{i3}\Delta LPM_{t-i}$$

$$+ \sum_{i=0}^{k-1} a_4 \Delta LY_{t-i} + gLM_{t-k} + dLPD_{t-k} + qLPM_{t-k} - JLY_{t-k} + n_t \quad . \qquad (10)$$

This is then followed by a testing-down procedure in search of a suitably parsimonious final model. In contrast with the previous model, both long run and short-run elasticities are estimated together in this approach. In our model

we tested k=1 against k=2 and obtained F(4,3) = 0.872 [p = 0.568]. Clearly, the restrictions were not rejected by the data. Therefore, the first ECM was preferred and the estimated model is

$$\Delta LM_{it} = -8.15 - 0.23\Delta LM_{t-1} + 0.80\Delta LPD_t - 0.44\Delta LPM_t + 1.12\Delta LY_t$$

(-3.62) (1.98) (1.32) (-0.90) (1.90)(1.16)

$$- 0.69\, LM_{t-1} + 1.17\, LPD_{t-1} - 0.43\, LPM_{t-1} - 1.99\, LY_{t-1} + \nu_t \text{, (11)}$$

(-3.94) (3.42) (-1.40) 3.24)

where R^2 = 0.80, normality test $\chi2(2)$ = 0.732 [p = 0.693], Durbin-Watson = 1.66, Serial Correlation LM $\chi2$ (1) = 2.48 [p = 0.115], heteroscedasticity LMARCH $\chi2(1)$ = 0.643 [p = 0.423].

Similar to those of Engle-Granger ECM, the statistics obtained from Hendry's ECM have passed all the diagnostic tests at the 5% significant level.

The long-run elasticities of consumer price, import price, and income are obtained by transforming the (t-1) lagged independent variables of equation (11) into the disequilibrium error form. By so doing, the *long-run* elasticities of consumer price, import price, and income can be calculated as follows:

$$\varepsilon_{Y}^{LR} = \delta / \gamma = 2.87, \qquad \varepsilon_{PD}^{LR} = \theta / \gamma = 1.68, \qquad \varepsilon_{PM}^{LR} = \vartheta / \gamma = -0.62,$$

while the short-run elasticities of those variables are given by the estimated coefficients on the zero lag terms:

$$\varepsilon_{Y}^{SR} = 1.12, \qquad \varepsilon_{PD}^{SR} = 0.80, \qquad \varepsilon_{PM}^{SR} = -0.44.$$

Furthermore, the coefficient of the disequilibrium error is 0.62, which is about the same as that derived from the Engle-Granger approach (0.617). Thus, both equations imply that roughly 62% of any disequilibrium between actual and equilibrium import demand in any one year is made up within the

next year. It is interesting to note that the coefficients estimated from Hendry's ECM do not deviate significantly from those of Engle-Grangers ECM. More importantly, both approaches lead to equations with approximately 62% of the long-run effects being felt after one year.

Finally, in order to discriminate the performance of the two models; we looked at their ex-post forecasting abilities over the period 1993-1997. Table (5) reports some commonly used measures of forecasting error over the post-sample period. It is clear from Table (5) that the Engle-Granger ECM outperforms Hendry's by producing smaller forecast errors, over this brief period. Hence, we concluded that Engle-Granger's ECM provided the best fit to the Saudi Arabia aggregate import demand.

Table (5): Measures of Forecasting Errors for the Two ECMs

	Mean Prediction Errors	SS Prediction Errors	Root MSE	Predicted Failure Test (F)
Engle-Granger	-0.008	0.063	0.127	0.853 [p=0.535]
Hendry	-0.222	0.061	0.248	1.542 [p=0.255]

Summary and Conclusion

The main objective of this study is to estimate and analyze the aggregate import demand function for Saudi Arabia. Cointegration and error correction techniques have been used to measure the impact of

economic activity (proxied by real GDP), real import prices, and real domestic prices on import demand.

The stationarity tests of the above four time series included in the study showed that they are non-stationary and integrated of degree one. The variables were found to be cointegrated and, therefore, possess a long-run equilibrium relation. The Johansen procedures for cointegration based on the VAR model have been used to find long and short-run estimates of the income and price elasticities of imports. Also, the error correction model (ECM) was used to specify the short-run dynamics for Saudi imports using two different approaches.

Both Engle-Granger’s type ECM and Hendry’s type ECM were used to analyze the import demand function for Saudi Arabia. It was found, in both models, that domestic price, import price and income are all-important in determining the aggregate import demand. The long and short run elasticities for the three determinants were also calculated as well as the coefficient of the disequilibrium error. The estimates obtained from Hendry's ECM did not deviate significantly from those of Engle-Granger’s. More importantly, both approaches lead to equations with approximately 62% of the long-run effects being felt after one year. The empirical results of the estimated cointegration vector show that Saudi aggregate import demand tends to be income elastic (2.60). This result implies that demand for imports in Saudi Arabia increases more proportionately to an increase in the level of the country’s economic activity (GDP). Moreover, the empirical results show that aggregate import demand tends to be elastic with respect to domestic prices while being inelastic with respect to import prices.

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دالة الطلب الكلية من الواردات في المملكة العربية السعودية: مدخل لتصحيح الخطأ

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ملخص

أدى ارتفاع الدخل النفطي في المملكة العربية السعودية إلى نمو وتغير هيكلي كبير في مختلف القطاعات الاقتصادية وفي مقدمتها قطاع التجارة الخارجية من صادرات وواردات. تهدف هذه الورقة إلى دراسة وتحليل دالة الطلب الكلية من الواردات في المملكة العربية السعودية من خلال تطبيق تحليل التكامل المشترك، وبناء نموذجين لتصحيح الخطأ أحدهما باستخدام طريقة أنجل وجرنجر ذات الخطوتين معتمدين على مدخل يوهانسون لبناء العلاقات طويلة الأجل ومن ثم تطبيق نموذج تصحيح الخطأ، والآخر باستخدام طريقة هندري. وقد تم تحديد المرونات الدخلية والسعرية طويلة وقصيرة الأجل واختبار قدرة النموذجان التنبؤية.