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Article

Short-Run and Long-Run Elasticities of Diesel Demand in Korea

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Abstract: This paper investigates the demand function for diesel in Korea covering the period 1986–2011. The short-run and long-run elasticities of diesel demand with respect to price and income are empirically examined using a co-integration and error-correction model. The short-run and long-run price elasticities are estimated to be −0.357 and −0.547, respectively. The short-run and long-run income elasticities are computed to be 1.589 and 1.478, respectively. Thus, diesel demand is relatively inelastic to price change and elastic to income change in both the short-run and long-run. Therefore, a demand-side management through raising the price of diesel will be ineffective and tightening the regulation of using diesel more efficiently appears to be more effective in Korea. The demand for diesel is expected to continuously increase as the economy grows.

Keywords: elasticity; diesel demand; Korea; co-integration; error-correction model

1. Introduction

The issues of global warming and greenhouse gas emission reduction take more interest in the diesel as a transportation fuel. Diesel engines have some advantages over gasoline engines, such as high efficiency, low flashing point, high torque and low trouble rate. Furthermore, the developments of diesel engine makes it possible to greatly reduce the emission of diesel particulate matter (PM) and

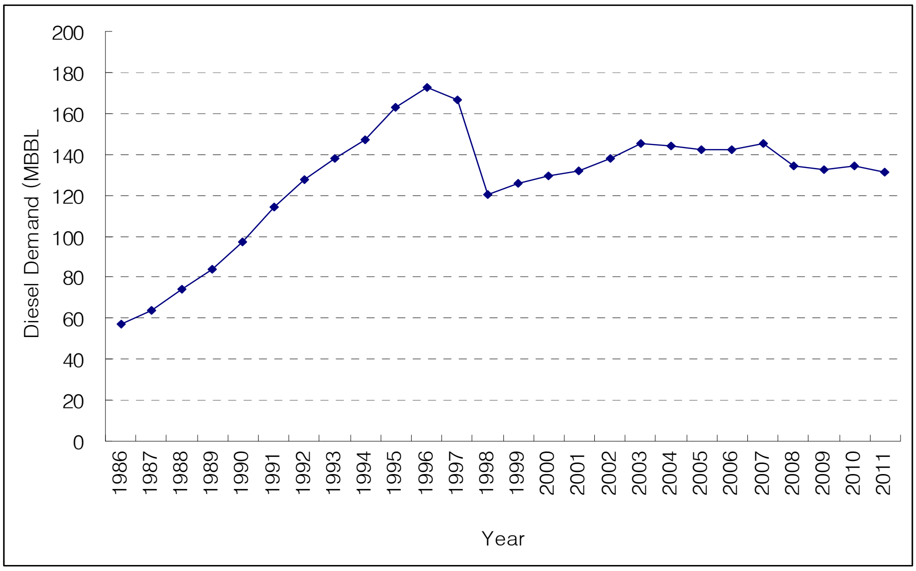
NO_x, regarded as shortcomings of diesel. Table 1 shows the polluting gas emission by each fuel [1]. Diesel engines shows higher official fuel efficiency than gasoline and liquefied petroleum gas (LPG) engines in the similar vehicle model.

Table 1. Polluting gas emission of diesel, gasoline and LPG.

Fuels	Engine	Polluting gas emission (g/km)					Official
(Vehicle model)	displacement	CO	HC	NO_X	PM	CO_2	fuel efficiency (km/L)
Diesel (i40 1.7)	1,685 cc	0.136	0.023	0.142	0.0001	149	18.0
Gasoline (sonata 2.0)	1,999 cc	0.111	0.014	0.014	-	159	14.7
LPG (sonata 2.0)	1,999 cc	0.340	0.007	0.011	-	169	10.5

In Korea, diesel has been used as a fuel for trucks, buses and heavy equipment used in construction. However, the share of diesel cars in the passenger car fleet has increased in recent years, since the common rail engine has introduced in 2003. Demand has increased 4.6% annually from 1986 to 2010 as shown in Figure 1 [2]. In particular, diesel demand in Korea from 1986 to 1997 has risen sharply up to 13.7% per annum. After 1998, diesel demand in Korea has rapidly decreased as the real economy slowed down due to the financial crisis. Since then, the diesel demand in Korea has maintained roughly constant.

Figure 1. Diesel demand in Korea.



The price for diesel has been increasing more sharply than that for gasoline since 1998. For example, the prices for diesel and gasoline in 1998 were 553 and 1120 Korean won per liter, respectively, but those in 2011 were 1778 and 1944 Korean won per liter, respectively. The price for diesel was 49.3% of that for gasoline in 1998, but 91.5% in 2011. Thus, the demand for oil products has been overall increasing, but the demand for diesel shows very little growth.

Understanding the sensitivity of diesel or gasoline demand to changes in prices and income has important implications for policies related to climate change, optimal taxation, pollution abatement and national energy security. Thus, a great deal of attention has been directed towards studying the demand in some studies [3]. Eltony [4,5] found low income and price elasticities in short-run and long-run by the co-integration and error-correction model (ECM). However, most studies do not provide a consensus on both short-run and long-run. Liddle [6,7] suggested that additional variables could be included (e.g., vehicle stock, vehicle-miles and motor fuel use, *etc.*), under the assumption that there are the systemic relationships among the variables. Table 2 contains the elasticities of transportation fuel (GCC consists of Bahrain, UAE, Oman, Saudi Arabia, Qatar and Kuwait. OECD is composed of Canada, the United States, Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, UK, Australia, Japan, and Turkey) [4,5,7–17].

Table 2. Literature review on estimates of the elasticities of gasoline and diesel demand.

Country	Transportation fuel	Period	Long-run elasticity		Short-run elasticity		Sources
			Income	Price	Income	Price	
Kuwait	Gasoline	1970–1989	0.92	−0.46	0.47	−0.37	Eltony [4]
GCC	Gasoline	1975–1993	0.48	−0.17	0.31	−0.11	Eltony [5]
OECD	Gasoline	1978–2005	0.20 0.34	−0.19 −0.43	0.28	−0.16	Liddle [7]
India	Gasoline	1972–1994	2.68	−0.32	1.18	−0.21	Ramanathan [8]
OECD	Gasoline	1960–1985	1.17	−0.79	0.41	−0.24	Graham *et al.* [9]
Brazil	Gasoline	1974–1994	0.12	−0.46	0.12	−0.09	Alves and Bueno [10]
Namibia	Diesel	1980–2002	2.08	−0.11	-	-	de Vita *et al.* [11]
Greece	Gasoline	1978–2003	0.79	−0.38	0.36	−0.10	Polemis [12]
	Diesel	1978–2003	1.18	−0.71	0.42	−0.07	
South Africa	Gasoline	1978–2005	0.36	−0.47	-	-	Akinboade *et al.* [13]
Fiji	Gasoline	1978–2005	0.43	−0.24	2.56	−0.36	Rao and Rao [14]
US	Gasoline	1976–2008	0.07	−0.25	-	-	Park et al [15]
Senegal	Gasoline	1970–2008	1.14	−3.01	0.46	−0.12	Sene [16]
Korea	Gasoline	1996–2010	1.40	−7.40	1.26	−3.81	Lim and Yoo [17]

The studies dealing with demand function can provide the information of consumer's demand behavior and economic policy's effects. Especially, the analysis on the income and price elasticity of diesel demand gives important implications toward reasonable tax policies. Economists have highlighted the role of fuel taxes, on the grounds of economic efficiency. However, their effectiveness might be limited in the short term, if the price elasticity of demand is low and the income is high. This is generally the case. The objective of this paper is to obtain both short-run and long-run price and income elasticities for diesel demand in Korea and to look into how the demand is elastic to price and income changes.

The remainder of the paper is organized as follows. Section 2 provides an overview of the methodology adopted here and explains the data employed. Section 3 presents the results. Some concluding remarks are made in the final section.

2. Methodology and Data

The econometric models considered in this paper for estimating the diesel demand are generally drawn from the demand studies discussed earlier. Dahl and Sterner [18] have found that income and price to be the explanatory variables in determining demand. Several studies analyzing diesel demand using co-integration and error correction have used the following model, which will be used in this study. The diesel demand function can be specified as follows:

$$D_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 P_t + u_t \tag{1}$$

where *D*, *Y* and *P* represent natural logarithms, respectively: *D* is the per capita diesel consumption; *Y*, the real per capita gross domestic product (GDP) that is seasonally adjusted quarterly data; *P*, the real diesel price; and u_t, the error term. We use quarterly time series data for the period 1986–2010. Korean Statistical Information Service (KOSIS) and Monthly Energy Statistics published by Korean Energy Economics Institute provide statistical data and diesel prices in Korea. The seasonally adjusted real GDP is measured in billion won as well as the population total. The prices for diesel have been deflated by the consumer price index (2005 = 100).

The main advantages of using this model are easy to distinguish between short-run and long-run. Sterner [3] have emphasized that the estimation of ECM involves three steps. The first step is to examine the stationarity of variables by unit root test. If the variables are found to be non-stationary, we must take the first difference and then examine the stationarity of differenced ones. Secondly, if they are found to be co-integrated, the long-run elasticities are estimated from the co-integration equation. Finally, the short-run elasticities can be estimated from an ECM.

The time series of variables is required to be stationary. It has been shown that estimating non-stationary data can yield spurious causality [19,20]. Therefore, following Engle and Granger [21], we first test the unit roots of the variables (*D*, *Y*, *P*) to confirm the stationarity of each one. This is often done by the augmented Dickey-Fuller (ADF) [22,23] and the Phillips-Perron (PP) [24] tests. In particular, the PP test is known to be robust for a variety of serial correlations and time-dependent heteroscedasticities. Thus, we employ the PP test to ensure the stability. If any variables are found to be non-stationary, we must take the first difference and then apply to estimate elasticities for diesel with differenced data.

According to Engle and Granger [21], the concept of co-integration can be defined as a systematic co-movement among two or more economic variables over the long run. If the two variables may have the property that a particular combination of them $Z = D - \alpha_1 Y - \alpha_2 P$ is stationary, then the coefficients α_1 and α_2 are long-run income and price elasticities, respectively. Akaike's information criterion (AIC) specifies the rule used to choose an optimal lag length (number of augmenting lags), which is described Pantula *et al.* [25]. The Johansen co-integration test procedure [26] is used for this purpose, since AIC is used for this test. AIC is a measure of the relative goodness-of-fit to describe the trade-off between bias and variance in model. Therefore, the lag length which minimize AIC can minimizes the estimated information loss.

If the variables are I(1) and co-integrated, co-integrating variables are expected to restore themselves to their long-run equilibrium whenever there is a drift, and ECM is constructed to estimate the short-run behavior [5]. The ECM can be estimated accordingly as follows:

$$\Delta D_t = \beta_{10} + \sum_{i=0}^{L_{11}} \beta_{11i} \Delta Y_{t-i} + \sum_{j=0}^{L_{12}} \beta_{12j} \Delta P_{t-j} + \sum_{k=1}^{L_{13}} \beta_{13k} \Delta D_{t-k} + \beta_{14} \varepsilon_{t-1} + u_{1t} \quad (2)$$

where Δ is the difference operator, L's are the numbers of lags, β's are parameters to be estimated, u_t's are the serially uncorrelated error terms, and ε_{t-1} is the error-correction term (ECT), which is derived from the long run co-integration relationship. Coefficients β_{110} and β_{120} give the short-run income and price elasticities, respectively, while β_{14} represents the speed of adjustment toward the long-run equilibrium [4]. In the ECM, the optimal lag lengths are chosen by using AIC suggested in Pantula *et al.* [25].

3. Results and Discussion

3.1. Results of Unit Roots and Co-Integration Tests

When testing for unit roots and co-integration, we have used a 0.05 threshold in this study. As discussed above, we employ PP test for the unit root test for the series of variables. The results are shown in Table 2. The p-values of PP values are smaller than 0.05. Thus, the null hypothesis of non-stationarity can be rejected for first differences of these series. All time-series are I (1).

The co-integration is to test whether a linear combination of each individually non-stationary time series is itself stationary. The results of the Johansen co-integration test for the series (D, Y, P) are reported in Table 3 (The numbers inside the brackets are the number of lags. The values in parentheses are p-values calculated under the null hypothesis of non-stationarity). The likelihood ratio tests show that the null hypothesis of absence of co-integrating relation (R = 0) can be rejected at 5% level of significance, and that the null hypothesis of existence of at most one co-integrating relation ($R \leq 1$) and two co-integration relation ($R \leq 2$) also cannot be rejected in Table 4 (The optimal lag length is chosen as four by using Akaike's information criterion described in Pantula *et al.* [25]. The p-values are calculated under the corresponding null hypothesis). This implies that there is one or more co-integrating equation at 5% level of significance. The optimal lag length is chosen as four by using Akaike's information criterion described in Pantula *et al.* [25].

Table 3. Results of Phillips-Perron unit root tests.

	Levels	First-differences
D	−20.954 [6] (0.058)	−108.355 [5] * (0.000)
Y	−10.658 [4] (0.393)	−80.052 [3] * (0.000)
P	−7.814 [3] (0.599)	−84.742 [2] * (0.000)

* Represent the rejection of the null hypothesis at the 5% level.

The co-integration vector equation is estimated to $Z = D - 1.4784Y + 0.5469P$ by using canonical co-integrating regression (CCR) suggested in Park [27]. Thus, the long-run income elasticity of diesel demand is 1.4784 and the long-run price elasticity is −0.5469. Both the two are statistically significant at the 1% level. The signs for the elasticity are alignment with economic theory, and diesel demand is elastic with respect to income, and is inelastic with respect to price. Furthermore, we can identify ECT, $\varepsilon_{t-1} = D_{t-1} - 1.4784Y_{t-1} + 0.5469P_{t-1}$, from estimated co-integration equation. In addition, we test Durbin-Watson statistic to detect the presence of autocorrelation in the residuals from the regression analysis. The estimated Durbin-Watson statistic give a value of 1.603 close to 2. Therefore, the CCR has no autocorrelation.

Table 4. Results of Johansen co-integration tests.

Null hypotheses	Likelihood ratio test statistic	*p*-values
H_0 ($R = 0$)	35.372 *	0.039
H_0 ($R \leq 1$)	17.609	0.060
H_0 ($R \leq 2$)	2.816	0.086

R denotes the number of co-integrating equation. * indicates the rejection of the null hypothesis at the 5% level.

3.2. Results of Error-Correction Model

In the ECM, the optimal lag lengths in Equation (2) are chosen by using AIC. The lag lengths of L_{11}, L_{12} and L_{13} in Equations (2) are chosen as 4, 3, and 4, respectively. The short-run price elasticity of diesel demand is estimated to be −0.3566 and the short-run income elasticity 1.5887. Both the two are statistically significant at the 1% level. The signs for elasticities are consistent with economic theory, and diesel demand is elastic with respect to income and in-elastic with respect to price.

In order to determine the most appropriate model, we conduct specification error tests. First, one important point to consider in estimating the ECM is that there might be a structural break. If there are any structural breaks, necessary adjustment to reflect the structural break should be made. We checked the model stability using CUSUM and CUSUMSQ tests suggested by Brown *et al.* [28]. The tests are appropriate for time series data and might be used if one is uncertain about when a structural break might have taken place. Moreover, the tests are quite general in that they do not require a prior specification of when the structural break takes place [29]. The null hypothesis is that the coefficients are the same in every period, *i.e.*, there is no structural break. Table 5 contains the results of the tests. Both tests suggest that the null hypothesis of absence of structural break cannot be rejected. Thus, the models are stable over time.

Second, it is important to consider whether there are any factors that are not included in this model but may affect the diesel demand. This aspect is a general specification error issue. Therefore, a fairly reasonable test to establish the specification error was conducted. It is interesting to determine whether there might be a significant bias in the estimates due to the omitted variables or non-exogeneity of the regressors. This is particularly important in this case since many scholars have tried to assess the effect on demand of a number of variables that are not included here because of the explicit focus on price and income.

Regression specification error test (called RESET) suggested by Ramey [30] was used in this study. In its simplest form, the procedure involves running a test regression of ΔD_t on the regressors included in Equation (2) along with the square of the predicted value of ΔD_t from the original regression. If the coefficient of the squared predicted-ΔD_t term in the test regression is not significant, the null hypothesis of no specification error can be maintained. The *t*-value associated with the squared predicted-ΔD_t term is 0.832, which is not statistically significant at any meaningful level. Thus, despite the simplicity of this model and the estimation procedure, there is no indication of any major specification problem.

Finally, this paper also used Durbin-Watson test statistic to detect the presence of autocorrelation in the residuals from the regression analysis. Furthermore, ECM include the lagged dependent variables, and therefore we employ the Durbin *h*- test. The estimated Durbin *h*-statistic give a value of −1.21 and *p*-value is 0.23. Therefore, the ECM model passes the test of autocorrelation.

Table 5 (The *p*-values are calculated under the null hypothesis of absence of structural break) gives the computed values of short-run and long-run elasticities for income and price. This implies that diesel demand is elastic with respect to income and the long-run elasticity is smaller than the short-run. This may be explained by rapid adjustment of consumer's automobiles stocks in the short run. The fact the long-run income elasticity was less than the short-run elasticity may suggest a problem with the model . This can be driven by omitted variable bias. For example, Liddle [6,7] considered vehicle stock in addition to income and fuel price. However, we cannot reject the null hypothesis of no specification error, as discussed above. Moreover, Rao and Rao [14] also found that the long-run elasticity is less than the short-run elasticity in Fiji. Alves and Bueno [10] detected that income elasticity does not differ from short-run and long-run in Brazil. This issue merits further investigation in the future. Furthermore, diesel demand is inelastic with respect to price and the long-run elasticity is larger than the short-run values in its absolute value. The rising prices will bring about a gradually decrease of diesel demand in both the short and long run. Finally, the coefficient of ECT, β_{14}, is significant (at the 1% level) and has a absolute value of 0.238 suggesting the diesel consumption adjusts toward is long-run level with almost 23.8% of the total adjustment occurring within the first quarter.

In conclusion, the estimated results of the elasticities of diesel demand are presented at Table 6. Comparing our results with the findings of Lim and Yoo's [17] study that deals with the gasoline demand function in Korea reveals that the diesel demand is less elastic to both price and income change than the gasoline demand.

Table 5. Results of CUSUM and CUSUMSQ tests.

	CUSUM test		**CUSUMSQ test**	
	Test statistic	***p*-values**	**Test statistic**	***p*-values**
Equation (2)	0.796	0.142	0.117	0.451

Table 6. Estimated results of the elasticities of diesel demand.

	Short-run	**Long-run**
Income elasticity	1.589 ** (5.12)	1.478 ** (17.39)
Price elasticity	−0.357 ** (−5.49)	−0.547 ** (−13.99)

** indicates statistical significance at the 1% level. The values in parentheses are *t*-values. The long-run elasticities are estimated by using co-integrating regression suggested in Park [27].

4. Conclusion

This study attempted to estimate the diesel demand Equation for Korea using co-integration and ECM. The results indicate that the growth in diesel demand can be attributed to rising incomes and falling prices during time period. The short-run and long-run price elasticities were computed to be

−0.357 and −0.547, respectively. The short-run and long-run income elasticities were estimated to be 1.589 and 1.478, respectively.

The diesel demand is inelastic with respect to price. We could reasonably expect that one percent increase in price gives rise to diesel consumption less than one percent. Furthermore, the estimated coefficients for price variables are significant and negative in their signs. These results are explained that rising price cause the consumers to use public transport rather than to utilize owned cars. Moreover, the long-run price elasticity is larger than the short run, indicating that the diesel demand response is more sensitive in the long run. This may be explained that the consumers' vehicles are constrained in the short run. However, in the long run, falling relative prices will introduce consumers to buy a bigger ones or another.

On the other hand, the diesel demand is elastic with respect to income. We could expect that one percent increase in real income gives rise to diesel consumption more than one percent. This may be explained by rapidly adjustment of consumer's automobiles stocks in the short run. Unlike price elasticities, the long-run income elasticity is smaller than the short run. This can be due to the type of the diesel vehicles. The diesel cars, like trucks and buses, make frequent use of the business. Moreover, most performance cars are based on gasoline engine in Korea. Moreover, when compared the price elasiticies, the consumers are more sensitive to rising income than price change. Therefore, the rich consumers travel more long-distance or use more vehicles for business transportation or private vacation, and buy more gasoline cars than diesel ones in the long run. Finally, the ECT is significant, suggesting that diesel consumption adjusts toward its long run level with about 23.8% of the adjustment taking place within the first quarter.

It is common that energy demand is in-elastic to price change and elastic to rising income. Our findings implicated that the tax policy is useless rather than direct demand management policy, because the consumers are not sensitive to price change. Since then, tightening the regulation of using diesel more efficiently will be more effective. In addition, the policy makers also need to interest on the limitation of price change, since diesel is mainly used for private travel as well as public transportation.

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