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An Empirical Analysis of Gasoline Demand in Mexico using Cointegration Techniques

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AN EMPIRICAL ANALYSIS OF GASOLINE DEMAND IN MEXICO

USING COINTEGRATION TECHNIQUES

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To my parents, Gloria and Gilberto Elizalde, along with my sister, Edith Elizalde, for the support and encouragement they have given me.

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Abstract

Economic expansion in Mexico has caused fuel consumption in Mexico to increase. Because Mexico does not have sufficient refinery capacity, over 40 percent of total gasoline consumed is imported. This has implications for the balance of payments. In this paper, gasoline demand is empirically examined using co-integration and error correction approaches. Results show no evidence of long-run equilibrium in the Mexican gasoline market. This is potentially attributable to the regulatory regime that governs energy markets in Mexico. Regulated price adjustments that are not consistent with prevailing market conditions run the risk of misallocating resources. Permitting greater flexibility in private gasoline retail markets may prove beneficial in Mexico.

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Chapter 1: Introduction

Demand for gasoline in Mexico has grown substantially. Between 2000 and 2006, gasoline demand increased 35 percent while diesel demand only increased 21 percent (Secretaría de Energía [SENER], 2008). Economic expansion accounts for the increase in fuel demand. As incomes have grown, increases in car ownership have also occurred.

Greater fuels consumption has attracted political attention because of its implications for the balance of payments. Refinery capacity in Mexico is insufficient to meet gasoline demand. Between 1990 and 2002, energy imports increased 124 percent from 108 thousand barrels per day (kbpd) to 243 kbpd (SENER, 2002). At present, gasoline imports are required to meet more than 40 percent of total gasoline demand (SENER, 2008). This has contributed to a large balance of trade deficit.

The objective of this research is to develop an econometric model that analyzes the behavior of gasoline consumption in Mexico. Income and price elasticities will be measured for the different product categories employed. Time series techniques will be utilized to obtain these estimates.

Remaining sections of the study are organized as follows. Chapter 2 provides a review of recent empirical studies on gasoline demand. Chapter 3 summarizes data and model specification. Chapter 4 reports empirical results. Chapter 5 summarizes principal results and policy implications. A data appendix is included at the end of the study with all historical data listed.

Chapter 2: Literature Review

Literature regarding the study of gasoline demand is extensive. Subsequent to the 1973 energy crisis, understanding the determinants of fuel demand has been of ongoing interest to economists. Studies cover a broad range of fuel demand and the dynamic relationships between consumption and economic stimuli.

Gasoline demand studies sometimes reach very different results. The high level of heterogeneity among the parameters that determine fuel demand can be explained by the model specification, type of data, geographical area of study, time period, functional form and econometric technique. The most commonly included variables in these models are real income and the real price of gasoline. Additional variables sometimes include the stock of vehicles and vehicle efficiency. Empirical analyses may employ individual household data or aggregate data for a region or a country. Sample data can be time series, cross-sections, or pooled cross-section-time series.

There are several time series estimation methods commonly utilized to analyze gasoline demand. These alternative methods may or may not yield similar results. Bhaskara Rao & Rao, (2009) compare several alternative time series methods employing data for Fiji from 1970-2005. The selected methods include fully modified ordinary least squares, maximum likelihood methods, and error correction approaches. Estimates of the long-run parameters are found to be similar for all selected methods. Price and income elasticities are about -0.20 and 0.45, respectively.

Given that elasticity estimates for gasoline demand vary greatly, discussion of individual results must be made on a clear understanding on the method employed and the empirical context for the estimation. However, Graham & Glaister (2002), by compiling the results of a large collection of international gasoline demand studies, find that there exists substantial consistency among results. With respect to price, typically, short-run elasticities are close to -0.3 and for the long-run range between -0.6 and -0.8. Short-run income elasticity is normally estimated in the range of 0.35 and 0.55, while for the long-run it typically falls between 1.1 and 1.3. Short-run and long-run effects of gasoline prices on traffic levels tend to be less pronounced than the impact of prices on gasoline demand. Therefore, it is safe to say motorists always find ways to economize fuel use once given time to adjust.

Previous work for Mexico has focused on establishing the impact on gasoline demand due to price differentials between Mexico regions and the United States. Many of the studies for other international economies explicitly deal with the non-stationary nature of time-series. Examples include Ramanathan (1999), Bentzen (1994), Alves & De Losso da Silveira Bueno (2003), Cheung & Thomson (2004), and Akinboade, Ziramba, & Kumo (2008). Such an effort has not previously been attempted for gasoline demand in Mexico.

Ramanathan (1999) examines the relationship between gasoline consumption, national income, and the price of gasoline for India using co-integration and error correction techniques. In the long run, the income elasticity is found to be quite high at 2.682 while the price elasticity is fairly low at -0.319. The low price elasticity can be explained by the fact that energy is necessary for the development of a low-income country such as India. Short-run income and

price elasticity are 1.178 and -0.209, respectively. Gasoline consumption adjusts towards its long-run equilibrium at a relatively slow rate, with only 28 percent of the adjustment taking place within the first year.

Bentzen (1994) confirms the existence of a co-integrating long-run relationship between gasoline consumption, the stock of vehicles, and the real price of gasoline in Denmark. The model specification is based on empirical evidence that shows that miles driven per vehicle remained stable throughout the period 1950-90, and on the assumption that the stock of vehicles is influenced by real income, population, and the real price of gasoline. With this background, the model is specified so that income only affects gasoline demand through the stock of vehicles. The short-run and long-run elasticities for vehicles per capita are in both cases close to one, 0.89 and 1.04 respectively. In the long-run, this can be explained by the fact that an increase in gasoline demand with an increase in the stock of vehicles will be offset by an increase in fuel efficiency. In the short-run, a change in the stock of vehicles must have a nearly equal effect on gasoline demand. The price elasticity in the short-run is -0.32 and -0.41 in the long-run.

Alves & De Losso da Silveira Bueno (2003) analyze the short-run and long-run behavior of gasoline demand in Brazil using co-integration techniques. Given that alcohol is utilized as an alternative automobile fuel in Brazil, its price is included as an additional variable permitting the estimation of the cross-price elasticity between the two substitutable goods. It is found that gasoline demand is inelastic to both income and price changes. As expected for substitutes, the cross price elasticity is positive. However, its absolute value is relatively low; 0.4803 in the

long-run and 0.2297 in the short-run. This potentially reflects the high costs associated with shifting from a gasoline engine to an alcohol fueled engine. Income elasticities are almost identical in the short- and long-run estimated at 0.1217 and 0.1216, respectively. Price elasticity is estimated at -0.4646 in the long-run and -0.0919 in the short-run.

Cheung & Thomson (2004), using co-integration techniques, find that demand for gasoline in China is relatively inelastic to price changes, both in the short-run and in the long-run. The long-run income elasticity is close to one meaning that the future growth of gasoline demand will approximate the growth rate of the economy. The pattern of the short-run income elasticity is quite different when compared to similar studies for other countries. Short-run income elasticity is larger than long-run income elasticity. This can be explained by the country's rapid economic growth and accompanying increases in disposable income. A variance decomposition analysis of a 10 year forecast shows that per capita GDP explains 21 percent of the variance while the price of gasoline explains only 16 percent of it.

Akinboade et al. (2008) develop an econometric model to explain the behavior of motor gasoline consumption in South Africa. This study employs the bounds test approach to co-integration to empirically analyze the long-run relationship of price and income in the gasoline demand function. It is found that gasoline demand is both income and price inelastic. The long-run income and price elasticities are 0.36 and -0.47. The low price elasticity is attributed to an unreliable and inefficient public transportation system in South Africa. Price increases do not discourage gasoline consumption.

Ibarra Salazar & Cervantes (2008) compare the price elasticity in the northern border and non-border regions of Mexico. The main hypothesis is that gasoline service stations located along the northern border face competition from their counterparts in the United States, causing gasoline demand to be more price sensitive that it is in the rest of the country. Several equations are estimated using price, income, and different combinations of other economic variables. The price elasticity for the Northern part of Mexico ranges from -0.67 to -1.57 while that of the rest of the country ranges from -0.15 to -1.06. The recognized disparity of the price elasticity of gasoline demand between the northern border region and the non-border region has important policy implications; it justifies the price gap between the north and south regions, and is essential for unraveling the differential impact of prices on fiscal revenues.

By using a co-integration analysis and an error correction model, this study seeks to explicitly model the non-stationary nature of the time-series data utilized. The analysis is for the period 1997-2007, employing monthly data. The use of this method not only allows distinguishing the short-run from the long-run gasoline demand elasticities, but also permits identifying the speeds of adjustment toward the long-run values.

Chapter 3: Methodology

In economics, most observed time series are generated by stochastic processes that are non-stationary. This kind of trended data can potentially cause major problems for empirical econometrics. Ignoring the non-stationary nature of the data, when performing regressions, may give rise to spurious relationships and diagnostic test statistics that are too high, and therefore, make regression results extremely hard to correctly evaluate. The concept of co-integration allows combining series generated by random walk processes into a single series which is stationary.

The models developed for this study are drawn from those discussed earlier. One equation is estimated for each motor gasoline grade available in Mexico; regular unleaded and premium. The co-integrating equations take the following functional form:

$$\log CU_t = \beta_0 + \beta_1 \log Y_t + \beta_2 \log PU_t + \varepsilon_t \tag{1}$$

$$\log CP_t = \alpha_0 + \alpha_1 \log Y_t + \alpha_2 \log PP_t + v_t \tag{2}$$

where CU_t and CP_t are gasoline consumption for regular unleaded and premium gasoline, respectively, in thousand barrels per month; Y_t, is the Industrial Production Index (IPI), used as a proxy for GDP; PU_t and PP_t are the real prices of gasoline in 2002 pesos per liter, for regular unleaded and premium gasoline, respectively; and, ε_t and v_t are the stochastic error terms. The stock of vehicles is included as another variable in other models, but is not utilized here as

the number of cars that use diesel, unleaded, or premium gasoline, is not known. Justification for the use of logarithms is discussed later in this chapter.

When variables follow random walk processes, running a regression of one against another can lead to spurious regressions, uninterpretable test statistics, and goodness of fit measures that are relatively high. This is explained by the fact that the shape of the empirical distribution will be far from similar to that of the theoretical Student-t distribution. Also, given that a random walk does not have finite variance, the Gauss-Markov theorem does not hold, and, ordinary least squares (OLS) does not yield a consistent parameter estimator.

However, even though these variables can wander extensively, a linear combination between them can be stationary. If x_t is a vector of economic variables, it will be in equilibrium when the following linear constraint is satisfied:

$$\gamma' x_t = 0$$

In this case, the components of the vector x_t are said to be co-integrated and γ is the co-integrating vector. According to Engle & Granger (1987), components of x_t are said to be co-integrated of order *d, b,* denoted $x_t \sim CI(d, b)$, if components of x_t have the same order of homogeneity *d,* conventionally denoted $x_t \sim I(d)$, and there exists a vector $\gamma\ (\neq 0)$ so that $z_t = \gamma' x_t \sim I(d-b), b > 0$. Therefore, before running any regression, it is critical to identify the order of integration of each variable.

The augmented Dickey-Fuller (ADF) procedure is used to test dependent, and independent, variables for unit roots in level and first or second differenced forms. The test is based on the estimation of the following equation:

$$\Delta Y_t = \delta \cdot Y_{t-1} + \sum_{i=1}^{k} \delta_i \cdot \Delta Y_{t-1} + \varepsilon_t \tag{3}$$

with an inspection of the Student t-ratio for δ. Under the null hypothesis that $\delta = 0$, Equation (3) is a random walk and the process generating Y_t is nonstationary. Rejection of the null hypothesis, in favor of the alternative hypothesis that $\delta < 0$, implies that the process generating Y_t is integrated of order zero and is stationary.

In order to test the null hypothesis, approximated critical values must be used. Under the null hypothesis, Equation (3) represents a regression of a $I(0)$ variable on a $I(1)$. In this case, the t-ratio does not have the familiar Student t-distribution. A number of authors have calculated critical values through simulation. In this analysis, each test is evaluated utilizing MacKinnon (1996) critical values.

The ADF test allows for possible autocorrelation in the error process ε_t. The addition of lagged left-hand-side variables as explanatory variables can secure white noise properties for ε_t. The number of augmentation terms (k) to include in the equation is determined by using the Akaike Information Criterion (AIC) in conjunction with the Schwarz Information Criterion (SIC). The AIC is computed as:

$$AIC = -\frac{2l}{T} + \frac{2k}{T} \tag{4}$$

and the SIC is computed as:

$$SIC = -\frac{2l}{T} + \frac{k\log T}{T} \tag{5}$$

where l is the log likelihood estimate. These statistics provide information that helps determine the choice of augmentation in (3) by penalizing the addition of right-hand side variables that diminish the power of the test.

According to the methodology described in Charemza & Deadman (1997) for co-integration analysis, the next step is to test the variables for co-integration. That requires running the OLS regression for the long-run equations (1) and (2). The ADF equation (3) with the estimated residuals for each of the co-integrating regressions is then applied:

$$\Delta\hat{\varepsilon}_t = \delta \cdot \hat{\varepsilon}_{t-1} + \sum_{i=1}^{k} \delta_i \cdot \Delta\hat{\varepsilon}_{t-1} + \varphi_t \tag{6}$$

The residuals from these regressions are tested for stationarity. The residuals from Equations (1) and (2) can be interpreted as deviations from the long run path. As long as these deviations are stationary, the variables are said to be co-integrated.

The use of logarithms in these models should be noted. According to Cryer (1986), most time series exhibit increased dispersion with an increased level of the series. In this case, the larger the level of the series, the more variation there is around that level. Also, as time goes

by, there may be stronger correlation between neighboring values. This implies that there may be long periods where deviations of the process are well away from the non-stationary mean. If the variance of a series is proportional to the level of the series, a logarithmic transformation will yield a series with an approximately constant variance.

If the possibility of co-integration is not rejected, OLS applied to Equations (1) and (2) does not lead to spurious regression results. The long-run elasticities for unleaded gasoline are given by:

$$\frac{\partial \log CU_t}{\partial \log Y_t} = \beta_1; \qquad \frac{\partial \log CU_t}{\partial \log PU_t} = \beta_2,$$

where β_1 and β_2 are the income elasticity and the price elasticity, respectively. The long-run elasticities for premium gasoline are given by:

$$\frac{\partial \log CP_t}{\partial \log Y_t} = \alpha_1; \qquad \frac{\partial \log CP_t}{\partial \log PP_t} = \alpha_2,$$

where α_1 and α_2 are the income elasticity and the price elasticity respectively.

If all variables from Equation (1) and (2) are $I(1)$, and deviations from their long-run path are $I(0)$, models in first differences with an error correction mechanism can be developed. The error correction models (ECM) are estimated using the following equations:

$$\Delta \log CU_t = \theta_0 + \theta_1 \Delta \log Y_t + \theta_2 \Delta \log PU_t + \theta_3 \, \hat{\varepsilon}_{t-1} + V_t \tag{7}$$

$$\Delta \log CP_t = \omega_0 + \omega_1 \Delta \log Y_t + \omega_2 \Delta \log PP_t + \omega_3 \hat{v}_{t-1} + U_t \qquad (8)$$

The ECM's are used to estimate the short-run behavior of gasoline consumption. The coefficients θ_1 and θ_2 are, respectively, the short-run income elasticity and the short-run price elasticity for unleaded gasoline consumption. The coefficients ω_1 and ω_2 are, respectively, the short-run income elasticity and the short-run price elasticity for premium gasoline consumption. In addition, θ_3 and ω_3 are interpreted as the speeds of adjustment for any shock leading to a deviation from the long-run consumption equilibria.

Solving for the error term in Equation (1) at time t-1 yields the following expression:

$$\varepsilon_{t-1} = \log CU_{t-1} - \beta_0 - \beta_1 \log Y_{t-1} - \beta_2 \log PU_{t-1} \qquad (9)$$

Substitution of (9) into Equation (7) yields the following error-correction equation:

$$\Delta \log CU_t = (\theta_0 - \theta_3\beta_0) + \theta_1 \Delta \log Y_t + \theta_2 \Delta \log PU_t + \theta_3 \log CU_{t-1} + \theta_3 \beta_1 \log Y_{t-1} + \theta_3\beta_2 \log PU_{t-1} + V_t \qquad (10)$$

Equation (10) can be reduced into the following expression:

$$\Delta \log CU_t = C_0 + C_1 \Delta \log Y_t + C_2 \Delta \log PU_t + C_3 \log CU_{t-1} + C_4 \log Y_{t-1} + C_5\beta_2 \log PU_{t-1} + V_t \qquad (11)$$

Expected coefficients signs for Equation (1) are $\beta_1 > 0$, and $\beta_2 < 0$. For Equation (7) the expected coefficients signs are $\theta_1 > 0$, $\theta_2 < 0$, and $\theta_3 < 0$. Accordingly, if they hypothesized economic relationship holds, parameter signs for Equation (11) will be $C_1 > 0, C_2 < 0, C_3 < 0, C_4 < 0,$ and $C_5 > 0.$

With the same approach, solving for the error term at time t-1 in Equation (2) yields:

$$v_{t-1} = \log CP_{t-1} - \alpha_0 - \alpha_1 \log Y_{t-1} - \alpha_2 \log PP_{t-1} \qquad (12)$$

Substitution of (12) into Equation (8) yields the error-correction equation:

$$\Delta \log CP_t = (\omega_0 - \omega_3 \alpha_0) + \omega_1 \Delta \log Y_t + \omega_2 \Delta \log PP_t + \omega_3 \log CP_{t-1} + \omega_3 \alpha_1 \log Y_{t-1} + \omega_3 \alpha_2 \log PP_{t-1} + U_t \quad (13)$$

Equation (13) is rewritten as:

$$\Delta \log CP_t = g_0 + g_1 \Delta \log Y_t + g_2 \Delta \log PP_t + g_3 \log CP_{t-1} + g_4 \log Y_{t-1} + g_5 \log PP_{t-1} + U_t \quad (14)$$

Expected coefficients signs for Equation (2) are $\alpha_1 > 0$, and $\alpha_2 < 0$. For Equation (8) the expected coefficients signs are $\omega_1 > 0$, $\omega_2 < 0$, and $\omega_3 < 0$. In combination, this implies parameter signs for Equation (14) will be $g_1 > 0, g_2 < 0, g_3 < 0, g_4 < 0,$ and $g_5 > 0.$

Direct estimation of Equations (11) and (14) implies a regression of variables with different orders of integration. Variables expressed in first-difference form are $I(0)$ and variables in levels are $I(1)$. However, if the non-stationary variables are co-integrated, this property makes both sides of the equation $I(0)$. In other words, the error term is stationary. The next section describes the data and summarizes empirical results.

Chapter 4: Data and Empirical Results

The sample data used for estimation include 132 monthly observations from January 1997 to December 2007. This information has been obtained from the National Statistics Institute in Mexico (INEGI). The dependent variable in the models estimated is the total volume of sales of each gasoline grade. The nominal price of each gasoline grade is a simple average between the price in the northern border region of Mexico and the price in non border regions as published by INEGI.

Given that the gross domestic product (GDP) is only available quarterly, the Industrial Production Index (IPI) which is available monthly, is used as a proxy for GDP in this study. The Consumer Price Index is employed to deflate the nominal gasoline prices. The base period is June 2002 (06/2002 = 100). The IPI covers four major divisions of economic activity in Mexico: electricity, water, and gas; mining; construction; and manufacturing. The IPI is an inflation-adjusted measure of industrial activity with 1993 as its base year (1993 = 100).

The tests results for stationarity of the variables are presented in Table 1. The table shows the t-statistics from the ADF test for the variables in level and first differenced form using the SIC to determine the lag structure. Both the SIC and, the AIC yield virtually the same results. All of the t-statistics for the variables in levels are statistically insignificant, which means that the null hypothesis of a unit root cannot be rejected. In contrast, the t-statistics for CP_t, CU_t, PP_t, and PU_t in their first differenced forms are statistically significant at the one-percent level. The t-statistic for Y_t in its first differenced form is statistically significant at the five-percent level. This implies that all of the variables are integrated of order one.

Consequently, co-integration models are estimated with undifferenced data, and the error correction models with first-differenced data.

Table 1: Augmented Dickey-Fuller Test for Stationarity

Variables	Levels	First differences
$\log(CU_t)$	4.431	-4.564**
$\log(CP_t)$	1.912	-4.737**
$\log(PU_t)$	-2.646	-12.542**
$\log(PP_t)$	-1.304	-13.016**
$\log(Y_t)$	-1.419	-2.942*

*Denotes significance at the 5 percent level.
** Denotes significance at the 1 percent level.

Regression output for the long-run consumption Equations (1) and (2) is reported in Table 2. The t-statistics appear in parentheses. Both models fail to provide solid evidence of co-integration. In both cases the low Durbin Watson (DW) statistic is a potential indication of non-stationary residuals. Inspection of the Q-statistic, which is used for a multi-period autocorrelation test, helps confirm this result. The critical value at the one-percent significance level is 158.95. As the Q-statistic for both models is greater than the critical value, the null of no co-integration cannot be rejected.

The apparent lack of co-integration implies there is no long-run equilibrium between the dependent and independent variables. Also unexpectedly, the price elasticity coefficients have positive signs. The long-run income elasticites for unleaded and premium gasoline are 1.27 and 1.07, respectively. These coefficients are close to those reported by other countries. Only the income coefficients appear to be statistically significant at the one-and five-percent criteria. However, the presence of serial correlation prevents meaningful inferences from being made about the estimated parameters.

Table 2: Results of Co-integrating Regressions

Variable	Regular	Premium
C	-0.4505	0.2324
	(-0.8464)	(0.4072)
LOG(PRICE)	0.3536	0.3943
	(1.5381)	(1.8038)
LOG(INCOME)	1.2722	1.0794
	(8.1149)**	(6.2489)**
R-squared	0.6203	0.5831
Adjusted R-squared	0.6144	0.5766
Standard error of regression	0.0897	0.0890
Sum of squared residuals	1.0382	1.0227
Log likelihood	132.4939	133.4815
F-statistic	105.3531	90.2118
Durbin-Watson statistic	0.4864	0.4357
Augmented Dickey-Fuller t-statistic of residual	-0.4748	-2.0871
Q-statistic of residual	727.98**	356**

** Denotes significance at the 1 percent level.

Estimation results for the traditional ECM procedure appear in Table 3. The short-run income elasticities have the correct sign for both models and satisfy the one-percent significance criterion. The short-run price elasticity only exhibits the correct algebraic sign for the unleaded gasoline consumption model. None of the short-run price elasticities are statistically significant at the five-percent level. The error correction terms represent the speed of adjustment toward the long-run equilibrium. They are estimated to be -0.1470 and -0.1721 for regular and premium consumption respectively, signifying 14.7 percent and 17.2 percent of the adjustments towards long-run equilibrium occur during the first month. Therefore, it takes regular and premium gasoline consumption 6.8 and 5.8 months respectively, for the disequilibrium in the prior period to be fully reversed.

The error correction mechanisms in both models exhibit the correct algebraic sign and are statistically significant. However, the autocorrelation test fails to provide evidence in favor of co-integration. Both Q-statistics are above the one-percent significance level critical value of 158.95. This means residuals are autocorrelated. This comes as no surprise given that the error correction terms refer to the residuals derived from the long-run co-integrating equations and failed to show stationarity properties.

Table 3: Traditional Error Correction Empirical Output

Variable	Regular	Premium
C	0.0032	0.0045
	(0.7803)	(1.0404)
D(LOG(INCOME))	0.3855	0.3931
	(3.6056)**	(3.5547)**
D(LOG(PRICE))	-0.2706	0.1414
	(-1.4612)	(0.4372)
RESIDUALS(-1)	-0.1470	-0.1721
	(-3.0470)**	(-3.4254)**
R-squared	0.1470	0.1439
Adjusted R-squared	0.1268	0.1237
Standard error of regression	0.0470	0.0491
Sum of squared residuals	0.2801	0.3057
Log likelihood	216.8081	211.0819
F-statistic	7.2954	7.1180
Durbin-Watson statistic	2.8149	2.7022
Augmented Dickey-Fuller t-statistic of residual	-0.3093	-1.5624
Q-statistic of residual	186.71**	168.59**

** Denotes significance at the 1 percent level.

Estimation results for the second set of error correction equations appear in Table 4. Similar to the traditional error correction approach, short-run income elasticities exhibit the correct algebraic sign and are statistically significant at the one-percent significance level. With respect to short-run price elasticity estimates, as was the case with the traditional approach,

only the unleaded gasoline consumption model exhibits the hypothesized algebraic sign. None of these estimates are statistically significant. Long-run income elasticities estimates for both models have algebraic signs that run counter to what is hypothesized. Long-run price elasticity estimates are with the correct algebraic sign but statistically insignificant.

The intercept term in Equation (11) is negative. Given that the model is estimated in first-differences, implies a decrease in consumption of unleaded gasoline. This runs counter to what is expected for gasoline consumption in the sample period under consideration. However, none of the associated t-statistics for the intercepts are statistically significant.

As with the traditional ECM approach, the residuals are also autocorrelated. Both Q-statistics lie above the one-percent significance level. This confirms disequilibrium in Mexico's gasoline market given that for an error correction representation to hold, co-integration is a necessary condition.

Empirical results suggest gasoline price adjustment magnitudes are different from those required to produce market equilibrium. This is no surprise given the relatively rigid prices set by the federal government. In Mexico, gasoline prices are determined by an international reference cost of production, transportation costs, and taxes. Under these circumstances, prices are insensitive to changes in the supply-and-demand balance.

According to Lajous (2009), government entities responsible for setting prices do not have the technical and financial resources to adequately perform regulatory functions. It is unlikely, under these circumstances, that the public sector will be able to administer gasoline markets efficiently in Mexico. Results obtained in this study indicate that is what occurred during the sample period in question.

Table 4: Generalized Error Correction Models Empirical Output

Variable	Regular	Premium
C	-0.2682	0.0538
	(-0.8574)	(0.1503)
D(LOG(INCOME))	0.4134	0.3978
	(3.4577)**	(3.2145)**
D(LOG(PRICE))	-0.2766	0.1086
	(-1.3921)	(0.3170)
LOG(CONSUMPTION(-1))	-0.1489	-0.1731
	(-3.0584)**	(-3.4119)**
LOG(INCOME(-1))	0.2371	0.2039
	(2.0178)*	(1.5997)
LOG(PRICE(-1))	0.0353	0.0175
	(0.2574)	(0.1285)
R-squared	0.1502	0.1456
Adjusted R-squared	0.1162	0.1114
Standard error of regression	0.0472	0.0494
Sum of squared residuals	0.2790	0.3051
Log likelihood	217.0568	211.2057
F-statistic	4.4198	4.2588
Durbin-Watson statistic	2.8317	2.7103
Q-statistic of residual	187.67**	170.61**

** Denotes significance at the 1 percent level.

Chapter 5: Summary and Conclusions

The demand for gasoline in Mexico is empirically examined through the use of co-integration and error-correction techniques. This methodology has proven useful as means of distinguishing between long- and short-run effects in gasoline demand studies. In this study, the two gasoline grades available in Mexico are analyzed separately. The sample period under consideration is from 1997 to 2007.

Gasoline demand is modeled as a function of price and income. Price elasticities are of great interest for the federal government. They help estimate the effect of gasoline prices in fiscal revenue collected through taxes applied to gasoline sales. Therefore, understanding gasoline demand responses to price and income changes is critical for structuring a policy that helps support the public sector budget.

Empirical results show no evidence of a co-integrating relationship between the variables. The two-step Engle and Granger procedure to construct an ECM yields inconsistent results. Under the assumption of equilibrium, a second approach was taken to construct an ECM. This involved direct estimation of the short- and long-run coefficients from the error-correction equation. As with the traditional approach, serially correlated residuals and opposite relationships between some of the variables to those hypothesized, point to a lack of long-run equilibrium in the Mexican gasoline market.

The regulatory price regime that exists in Mexico's gasoline market may be the source of market disequilibrium. Regulated price adjustments that are inconsistent with prevailing market conditions run the risk of misallocating resources. The latter may hamper economic

performance and retard development. Permitting greater flexibility in private gasoline retail markets may prove beneficial in Mexico.

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Appendix

Sample data.

Date	Gasoline consumption (kbpd)		Price North Region[1] (pesos per liter)		Price all other regions[1] (pesos per liter)		Industrial Production Index	CPI
	Unleaded	Premium	Unleaded	Premium	Unleaded	Premium		
1997/01	486	334	2.79	3.1	2.91	3.2	109.18	56.94
1997/02	490	345	2.82	3.2	2.94	3.2	106.55	57.90
1997/03	483	349	2.85	3.2	2.97	3.3	109.26	58.62
1997/04	495	366	2.88	3.2	3.00	3.3	117.03	59.25
1997/05	498	383	2.91	3.3	3.04	3.4	116.37	59.79
1997/06	487	396	2.94	3.3	3.07	3.4	117.74	60.32
1997/07	508	441	2.97	3.3	3.10	3.4	118.60	60.85
1997/08	485	439	3.00	3.4	3.13	3.5	117.99	61.39
1997/09	498	466	3.03	3.4	3.16	3.5	117.80	62.15
1997/10	505	484	3.06	3.4	3.19	3.5	125.80	62.65
1997/11	484	473	3.09	3.4	3.23	3.5	119.25	63.35
1997/12	554	551	3.25	3.6	3.39	3.6	119.01	64.24
1998/01	500	500	3.27	3.6	3.41	3.7	117.02	65.64
1998/02	508	508	3.30	3.6	3.44	3.8	115.12	66.79
1998/03	504	504	3.32	3.7	3.46	3.8	127.26	67.57
1998/04	517	517	3.34	3.7	3.49	3.8	119.88	68.20
1998/05	494	494	3.37	3.7	3.51	3.9	123.51	68.74
1998/06	521	521	3.39	3.7	3.53	3.9	126.30	69.56
1998/07	519	519	3.41	3.8	3.56	3.9	125.66	70.23
1998/08	498	498	3.44	3.8	3.58	3.9	126.06	70.90
1998/09	507	507	3.46	3.8	3.61	4.0	125.83	72.05
1998/10	518	518	3.48	3.83	3.63	3.98	129.15	73.09
1998/11	498	498	4.04	4.44	4.21	4.63	123.78	74.38
1998/12	555	555	4.07	4.48	4.25	4.66	123.15	76.19
1999/01	479	446	4.1	4.52	4.29	4.72	119.81	78.12
1999/02	508	471	4.15	4.57	4.34	4.78	118.82	79.17
1999/03	531	490	4.19	4.61	4.38	4.82	130.64	79.90
1999/04	503	463	4.23	4.66	4.42	4.87	126.19	80.64
1999/05	499	459	4.28	4.71	4.47	4.92	128.83	81.12
1999/06	517	473	4.31	4.72	4.51	4.97	134.36	81.66
1999/07	507	464	4.36	4.8	4.56	5.02	132.65	82.20
1999/08	490	447	4.4	4.84	4.6	5.06	133.12	82.66
1999/09	511	465	4.45	4.9	4.65	5.12	131.69	83.46
1999/10	490	447	4.5	4.95	4.7	5.17	133.05	83.99
1999/11	524	476	4.53	4.99	4.74	5.22	132.41	84.73
1999/12	574	517	4.58	5.04	4.79	5.27	130.59	85.58

[1]A simple average between these two prices is used as nominal price for each gasoline grade.

Sample data (cont.).

Date	Gasoline consumption (kbpd)		Price North Region[1] (pesos per liter)		Price all other regions[1] (pesos per liter)		Industrial Production Index	CPI
	Unleaded	Premium	Unleaded	Premium	Unleaded	Premium		
2000/01	499	450	4.62	5.09	4.83	5.32	128.85	86.73
2000/02	527	474	4.66	5.13	4.87	5.36	129.99	87.50
2000/03	537	481	4.7	5.17	4.91	5.41	141.25	87.98
2000/04	510	456	4.73	5.21	4.95	5.45	131.13	88.48
2000/05	539	480	4.76	5.24	4.98	5.48	140.55	88.82
2000/06	537	478	4.8	5.29	5.02	5.53	143.82	89.34
2000/07	518	459	4.84	5.33	5.06	5.57	139.96	89.69
2000/08	541	478	4.89	5.48	5.11	5.73	143.66	90.18
2000/09	527	468	4.93	5.53	5.15	5.77	139.03	90.84
2000/10	526	465	4.96	5.56	5.19	5.82	142.02	91.47
2000/11	546	481	5	5.6	5.23	5.86	137.73	92.25
2000/12	568	498	5.04	5.65	5.27	5.91	128.90	93.25
2001/01	537	471	5.07	5.7	5.30	5.9	131.17	93.77
2001/02	540	474	5.10	5.7	5.33	6.0	125.46	93.70
2001/03	552	483	5.12	5.7	5.35	6.0	137.70	94.30
2001/04	534	465	5.15	5.8	5.38	6.0	127.11	94.77
2001/05	561	487	5.17	5.8	5.41	6.1	136.09	94.99
2001/06	556	483	5.20	5.8	5.44	6.1	137.23	95.21
2001/07	542	468	5.23	5.9	5.47	6.1	134.01	94.97
2001/08	556	478	5.26	5.9	5.50	6.2	137.08	95.53
2001/09	523	451	5.28	5.9	5.52	6.2	131.23	96.42
2001/10	561	482	5.31	6.0	5.55	6.2	135.73	96.85
2001/11	558	480	5.34	6.0	5.58	6.3	132.38	97.22
2001/12	582	496	5.37	6.0	5.61	6.3	124.44	97.35
2002/01	554	472	5.39	6.0	5.63	6.3	126.84	98.25
2002/02	549	469	5.40	6.1	5.65	6.3	122.52	98.19
2002/03	544	462	5.42	6.1	5.67	6.4	127.44	98.69
2002/04	562	476	5.44	6.1	5.69	6.4	137.97	99.23
2002/05	575	485	5.46	6.1	5.71	6.4	136.97	99.43
2002/06	540	456	5.49	6.2	5.74	6.4	135.74	99.92
2002/07	573	480	5.51	6.2	5.76	6.5	136.74	100.20
2002/08	570	477	5.53	6.2	5.78	6.5	137.26	100.59
2002/09	541	454	5.55	6.2	5.80	6.5	130.90	101.19
2002/10	575	482	5.57	6.24	5.82	6.52	137.82	101.64
2002/11	571	480	5.59	6.27	5.84	6.55	131.43	102.46
2002/12	624	524	4.44	6.29	5.86	6.57	126.53	102.90

[1]A simple average between these two prices is used as nominal price for each gasoline grade.

Sample data (cont.).

Date	Gasoline consumption (kbpd)		Price North Region[1] (pesos per liter)		Price all other regions[1] (pesos per liter)		Industrial Production Index	CPI
	Unleaded	Premium	Unleaded	Premium	Unleaded	Premium		
2003/01	581	488	4.63	6.29	5.87	6.58	128.46	103.32
2003/02	589	496	5.27	6.31	5.89	6.6	124.63	103.61
2003/03	579	485	5.64	6.32	5.9	6.61	133.46	104.26
2003/04	602	501	5.66	6.34	5.92	6.64	131.34	104.44
2003/05	607	506	5.19	6.36	5.93	6.65	134.32	104.10
2003/06	586	489	5.27	6.38	5.95	6.67	134.66	104.19
2003/07	610	505	4.9	6.39	5.96	6.68	134.82	104.34
2003/08	582	484	5.72	6.41	5.98	6.7	133.31	104.65
2003/09	596	496	5.73	6.42	5.99	6.71	131.42	105.28
2003/10	607	506	5.21	6.44	6.01	6.74	138.26	105.66
2003/11	590	491	5.15	6.46	6.02	6.75	130.19	106.54
2003/12	673	555	5.12	6.48	6.04	6.77	130.46	107.00
2004/01	609	502	5.04	6.66	6.05	6.96	129.79	107.66
2004/02	602	498	5.7	6.69	6.07	6.99	128.46	108.31
2004/03	640	527	5.82	6.7	6.08	7	142.51	108.67
2004/04	632	516	5.83	6.71	6.1	7.02	136.30	108.84
2004/05	619	507	5.84	6.72	6.11	7.03	137.64	108.56
2004/06	645	532	5.86	6.92	6.13	7.24	141.98	108.74
2004/07	640	528	5.88	6.94	6.15	7.26	139.53	109.02
2004/08	620	513	5.89	6.95	6.16	7.27	140.75	109.70
2004/09	633	526	5.91	6.98	6.18	7.3	139.33	110.60
2004/10	621	518	5.92	6.99	6.19	7.31	140.89	111.37
2004/11	655	547	5.94	7.01	6.21	7.33	137.84	112.32
2004/12	713	591	5.95	7.03	6.22	7.34	136.60	112.55
2005/01	626	520	5.96	7.1	6.24	7.4	132.88	112.55
2005/02	651	543	5.98	7.1	6.25	7.4	130.97	112.93
2005/03	675	560	6.00	7.1	6.27	7.4	136.12	113.44
2005/04	659	550	6.01	7.1	6.28	7.4	143.96	113.84
2005/05	657	548	6.03	7.1	6.30	7.4	142.26	113.56
2005/06	679	564	6.04	7.1	6.31	7.5	143.46	113.45
2005/07	646	537	6.05	7.1	6.33	7.5	138.47	113.89
2005/08	680	566	6.06	7.2	6.34	7.5	143.63	114.03
2005/09	684	572	6.09	7.2	6.37	7.5	141.27	114.48
2005/10	655	550	6.12	7.2	6.40	7.6	144.75	114.77
2005/11	697	584	6.14	7.3	6.42	7.6	142.25	115.59
2005/12	746	621	6.19	7.3	6.47	7.6	140.31	116.30

[1]A simple average between these two prices is used as nominal price for each gasoline grade.

Sample data (cont.).

Date	Gasoline consumption (kbpd)		Price North Region[1] (pesos per liter)		Price all other regions[1] (pesos per liter)		Industrial Production Index	CPI
	Unleaded	Premium	Unleaded	Premium	Unleaded	Premium		
2006/01	680	567	7.20	7.3	6.49	7.7	141.76	116.98
2006/02	699	584	7.15	7.4	6.51	7.7	137.65	117.16
2006/03	718	598	7.73	7.4	6.53	7.7	149.66	117.31
2006/04	690	575	8.27	7.4	6.56	7.8	142.45	117.48
2006/05	728	608	7.41	7.4	6.58	7.8	150.73	116.96
2006/06	725	607	7.41	7.5	6.60	7.8	153.96	117.06
2006/07	701	586	7.41	7.5	6.62	7.8	146.69	117.38
2006/08	732	610	7.41	7.5	6.64	7.8	151.29	117.98
2006/09	720	605	7.41	7.5	6.66	7.9	147.88	119.17
2006/10	709	596	6.84	7.54	6.68	7.89	151.10	119.69
2006/11	739	625	7.41	7.9	6.71	8.25	148.94	120.32
2006/12	777	661	7.41	7.92	6.74	8.29	142.48	121.02
2007/01	725	620	7.28	7.96	6.76	8.31	144.11	121.64
2007/02	736	634	7.41	7.99	6.79	8.35	138.20	121.98
2007/03	764	658	7.41	8.02	6.82	8.38	150.44	122.24
2007/04	727	626	7.41	8.05	6.85	8.42	145.46	122.17
2007/05	771	666	7.41	8.09	6.88	8.45	152.27	121.58
2007/06	773	670	7.41	8.11	6.9	8.48	154.20	121.72
2007/07	746	645	7.41	8.15	6.94	8.52	149.74	122.24
2007/08	769	666	7.41	8.19	6.97	8.56	155.45	122.74
2007/09	732	639	7.41	8.23	7.01	8.6	148.01	123.69
2007/10	779	681	7.41	8.27	7.01	8.65	156.34	124.17
2007/11	778	682	7.41	8.31	7.01	8.69	149.97	125.05
2007/12	820	718	7.41	8.35	7.01	8.73	144.34	125.56

[1]A simple average between these two prices is used as nominal price for each gasoline grade.

Vita

Mario Elizalde graduated from Instituto Tecnológico y de Estudios Superiores de Monterrey High School in Ciudad Juárez. He then attended The University of Texas at El Paso, where he received a Bachelor of Science in Civil Engineering in May 2008. He started his pursuit of a Master's degree in Economics, also at The University of Texas at El Paso, in summer 2008. During his studies, he was awarded the James Foundation Scholarship. Mario began working as a research assistant for the Center of Institutional Evaluation, Research and Planning (CIERP) while attending graduate school and currently is employed at CIERP full time.

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