

The cross-border effect on gasoline demand in Baja California

El efecto de la frontera en la demanda de gasolina en Baja California

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Abstract

This paper estimates gasoline demand in the border state of Baja California to determine the magnitude of the border effect through price elasticity relative to gasoline prices in California. A monthly database from January 1997 to December 2015 was used and the distributed lag autoregressive model was applied to estimate demand. It was found that, in the long run, gasoline demand in Baja California is approximately unitary with respect to the price in California. This confirms the importance of the cross-border effect on gasoline demand and provides evidence supporting fiscal policy towards this sector in that region. Keywords: gasoline demand, cross-border effect, gasoline tourism, autoregressive distributed lag model (ARDL), Baja California.

Resumen

En este artículo se estimó la demanda de gasolina del estado fronterizo de Baja California para determinar la magnitud del efecto frontera a través de la elasticidad con respecto al precio de la gasolina de California. Se utilizó una base de datos mensual de enero de 1997 a diciembre de 2015 y se aplicó el modelo autorregresivo de rezagos distribuidos para estimar la demanda. Se encontró que en el largo plazo la demanda de gasolina en Baja California es aproximadamente unitaria en relación con el precio de la gasolina en California. Esto confirma la importancia del efecto frontera en la demanda de gasolina y ofrece evidencia que justifica la política fiscal hacia ese sector en esa región.

Palabras clave: demanda de gasolina, efecto frontera, turismo gasolina, modelo autorregresivo de rezagos distribuidos (ARDL), Baja California.

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Introduction

The assessment by the Northern Border Free Zone Program (Secretaría de Economía, 2021) acknowledges the existence of price imbalances that influence economic activity in Mexico's border municipalities. Specifically, it notes that gasoline prices at gas stations on the U.S. side of the border were lower than in Mexico. Indeed, in December 2018, the assessment indicates that the price per liter of gasoline in Texas was \$10.62, while in Tamaulipas it was \$16.48, a difference of 55%. In response, the gasoline sector was included in the tax incentive program for the border region, known as the Northern Border Free Zone Program, which took effect on January 1, 2019.

The Ministry of Finance and Public Credit is responsible for setting gasoline subsidies, which are published weekly in the *Diario Oficial de la Federación*. Originally, the gasoline subsidy at the northern border was established based on the decree published on December 27, 2016, and updated on December 28, 2018, which includes reductions to the IEPS (Spanish acronym for *impuesto especial sobre producción y servicios*, special tax on production and services) applied to the price of gasoline.

Prior to the fiscal policy implemented in 2019, the border gasoline sector had already been subject to various regulatory and fiscal measures (Ayala Gaytán & Gutiérrez González, 2004; Haro López & Ibarrola Pérez, 1999; Ibarra Salazar & Sotres Cervantes, 2022, 2008). Until 2015, when the federal government set gasoline prices in Mexico, not only were fiscal policies different between the Mexican and U.S. markets, but so were the regulatory measures affecting the sector. Furthermore, changes in the peso (MXN)-dollar (USD) exchange rate were also a factor to consider in the gasoline price differential on both sides of the border. Now that gasoline prices in Mexico have been deregulated since 2017, the main factors influencing the price differential are the taxes levied by local and federal authorities, as well as the MXN-USD exchange rate.

The special consideration given to the border gasoline market in fiscal and regulatory policy stems from the influence of the price differential between gasoline in Mexico and the United States on gasoline demand in that region. This influence is referred to as the "border effect". By applying the model proposed in Banfi et al. (2005), this study aims to estimate the border effect on gasoline demand in the Baja California (BC) border region. A database consisting of a monthly time series from January 1997 to December 2015 is used to estimate the elasticity of gasoline demand in the BC border region with respect to the Mexican price (direct price elasticity) and the U.S. price (cross-price elasticity).

This work contributes to the literature on cross-border shopping—and particularly the so-called gasoline tourism—by examining the demand for gasoline in a Mexican border state. From a fiscal policy perspective, it will be important to determine how sensitive the demand for gasoline in BC is to gasoline prices along Mexico's northern border and the U.S. side of the border. Estimating these elasticities is essential for predicting the effect that changes in gasoline prices in Mexico (direct price elasticity) and in California, United States (cross-price elasticity) may have on demand. To illustrate, the more price-elastic the demand is, the greater the demand reduction will be

in response to a tax increase, assuming such increase is carried over to the retail price. On the other hand, a reduction in the price of gasoline in California would lead to a larger percentage decrease in gasoline demand in BC if the cross-price elasticity is greater than one in absolute value.

The paper is organized as follows: the next section presents the review of the literature, followed by an explanation of the methodology, then the results, and finally the conclusions.

Related literature

Cross-border effects on shopping have been documented for various goods. For example, in the study by Leal et al. (2010), goods affected by this phenomenon include alcoholic beverages, tobacco, gasoline and lottery tickets. The results of that review indicate that theoretical studies are consistent with empirical studies: tax differentials and transportation costs from neighboring territories influence consumers' purchasing decisions. According to Leal et al. (2010), this topic has been studied since the 1930s.

When it comes to borders within the same country, this phenomenon arises mainly from differences in taxes set by subnational governments. In the case of international borders, where different currencies are used for trade and market regulations differ, in addition to differences in fiscal policy, one would expect exchange rate fluctuations and differing market regulations to have an impact.

The literature related to this paper includes empirical studies that estimate gasoline demand at international and domestic borders to quantify gasoline tourism to some extent.

Banfi et al. (2005) use a panel of annual data from three Swiss border regions (Italy, France and Germany) to estimate demand using the generalized least squares method. Based on neoclassical consumer theory and a staged budgeting model, they define an empirical equation to estimate the cross-price elasticity of gasoline demand. In the various models estimated, they find that the cross-price elasticity ranges from 0.30 to 0.58, confirming that when gasoline prices rise at gas stations in neighboring countries, demand for gasoline along the Swiss border increases.

Leal et al. (2009) examine how differences in gasoline prices between neighboring autonomous communities in Spain influence fuel demand in one of them (Aragón). They use monthly data from January 2001 to March 2007 to estimate an error correction model. Their empirical model is based on the work of Kanbur and Keen (1993), who characterize consumer decision-making in neighboring regions using a partial equilibrium model that compares reservation prices adjusted for taxes and transportation costs. They find evidence that relative prices in neighboring regions are directly related to fuel demand in Aragón.

Building on the study by Leal et al. (2009), Kubovič (2013) uses monthly data from January 2005 to December 2012 to estimate gasoline demand in the Czech Republic and applying the error correction model and to analyze the correlation between demand and gasoline prices in neighboring countries (Germany, Poland, Austria and Slovakia). The results show that only in the case of Germany is there evidence of a direct correlation: the cross-price elasticity of Czech gasoline demand is 0.51 with respect to the price of German gasoline.

Kennedy et al. (2017) use an unbalanced panel of 543 gas stations from April 2013 to March 2015 to study cross-border fuel demand in Northern Ireland using a random-effects model. In their study, they considered market size, proximity to highways, local competition and service station characteristics. Their results show that gas stations near the border have average sales of 54.4% higher for diesel and 14.6% higher for gasoline than those located further from the border.

In Iran, rather than “gasoline tourism”, it has been recognized that the large price differentials with neighboring countries (Iraq, Turkey, Armenia, Azerbaijan, Turkmenistan, Afghanistan and Pakistan) lead to smuggling, which, according to Ghoddusi et al. (2018), involves the illegal transport of gasoline or the filling of large trucks or vans to transport it across the Iranian border. They analyzed the demand for gasoline in Iran based on price differentials and distance from neighboring countries. They used monthly data from 160 distribution agencies, from 2005 to 2014, to estimate demand using fixed effects. Their results show that the ratio of the price of Iranian gasoline to the price in neighboring countries makes the demand for smuggled gasoline more elastic.

Jansen and Jonker (2018) used microdata from a sample of Dutch consumers (35 766) with transactions between September 2013 and June 2015. Unlike other studies on gasoline tourism, their estimates using panel data techniques show evidence that the demand for gasoline among Dutch consumers living near the border is highly inelastic with respect to the cross-border price differential.

In this case, the dependent variable is a binary variable that indicates whether a consumer made a payment at a Dutch gas station on a given day. The independent variables include socioeconomic and demographic characteristics, price differentials relative to neighboring countries (Belgium and Germany), as well as variables reflecting the consumer’s distance from the border. The results show that the probability that a consumer on the Netherlands-Belgium border will purchase gasoline does not vary with distance to a gas station or with the price differential.

This result differs from that observed among consumers along the Netherlands-Germany border. In that case, the probability that a consumer living up to 10 km from the border will purchase gasoline increases by 1.1% if gasoline in Germany is 1% more expensive than in the Netherlands. This was apparently a temporary effect caused by a change in gasoline taxes. Their main conclusion is that price differentials do not appear to influence gasoline tourism among Dutch consumers significantly.

The study by Moshiri (2020), like that by Ghoddusi et al. (2019), also examines differences in the price elasticity of demand across different regions of Iran. They applied the AIDS (almost ideal demand system) model and consumer expenditure data from 2005 to 2016. They found that demand in border provinces is more elastic than in non-border provinces.

There are also studies on gasoline demand in Mexico’s northern border region. Haro López and Ibarrola Pérez (1999) used monthly data from January 1995 to July

1999 to estimate the price elasticities of demand in various border areas.¹ They found that the elasticity varies between -0.112 and -0.639 across the different empirical specifications and border regions.

Ayala Gaytán and Gutiérrez González (2004) estimated gasoline demand in the northern border region also using monthly data from January 1993 to May 2001. Their approach to calculate gasoline tourism was to eliminate the gasoline price differential between Mexico and various regions of the southern United States. Using this approach, they estimated that uniform prices would result in a 36% increase in gasoline sales.

Unlike the studies cited on Mexico, which include the border region, the study by Fullerton et al. (2012) focuses on gasoline consumption patterns in Ciudad Juárez, Chihuahua. It uses monthly data from January 2000 through December 2009, and estimates demand using the autoregressive integrated moving average (ARIMA) model. They estimate that the price elasticity of demand is 0.26.

The approach adopted by Ibarra Salazar and Sotres Cervantes (2008, 2022) was to compare the price elasticity of gasoline demand domestically with that in Mexico's northern border region. The working hypothesis, as in Ghoddusi et al. (2018) and Moshiri (2020), is that demand at border gas stations is more elastic than at non-border gas stations. This hypothesis is confirmed by estimating various fixed-effects models using a monthly panel dataset for Mexican states.

As noted, three different approaches have been used in the literature to quantify gasoline-driven tourism: comparing the price elasticity of demand in the border region with that in the rest of the country; comparing demand at actual prices with demand at the same price in both areas; and estimating the price elasticity of demand in one region relative to the price in the adjacent region.

This study applies the third approach by using a monthly time series from a specific border region in Mexico.

Model specification

The model developed by Banfi et al. (2005), which is applied in this paper, is based on the work of Becker (1965), Deaton and Muellbauer (1980) and Lin et al. (1984). The household production model (Becker, 1965) is adapted to transportation production

¹ Zone IA comprises the municipalities of Tijuana, Rosarito and Tecate in Baja California; Zone IB includes only Mexicali; Zone II comprises the municipalities of Nogales, Cananea, Naco, Puerto Peñasco, Plutarco Elías Calles, Caborca, Altar, Sáric and Agua Prieta in Sonora; Zone III covers the municipalities of Janos, Ascensión, Ciudad Juárez, Práxedes Guerrero, Guadalupe, Coyame, Ojinaga and Benavides in Chihuahua; Zone IV includes the municipalities of Ocampo, Acuña, Jiménez, Zaragoza, Piedras Negras, Nava, Guerrero and Hidalgo in Coahuila, as well as the municipalities of Anáhuac in Nuevo León and Nuevo Laredo in Tamaulipas; and Zone V encompasses the municipalities of Guerrero, Ciudad Mier, Miguel Alemán, Camargo, Gustavo Díaz Ordaz, Río Bravo, Valle Hermoso, Matamoros and Reynosa, in Tamaulipas.

and gasoline consumption. According to Banfi et al. (2005), consumer preferences are defined by the following utility function:

$$U = u(S(G, C, T), X; D, R), \quad (1)$$

where S represents transportation services that use gasoline (G), the family's car stock (C), and travel time (T) as inputs. The variable X represents the numerarie, while D and R represent demographic and geographic characteristics, respectively, that influence consumer's utility. As a result of a two-stage optimization process (Deaton & Muelbauer, 1980), it is possible to derive the gasoline demand:

$$G = G(P, I, C, D, R). \quad (2)$$

In (2), P represents the price of gasoline, and I represents household income. To adapt this general model to the aggregate demand for gasoline in a border region, it would be necessary to consider prices on both sides of the border, given that regulatory and fiscal frameworks as well as market structures can lead to different prices; and to also consider income and demographic factors on both sides of the border.

Thus, by applying the specification in Banfi et al. (2005) to the area of interest in this study, the following are included as explanatory variables: the price of gasoline at the BC border (P^{BC}); the price of foreign gasoline relative to the price of local gasoline (P^{CA}/P^{BC}); and an interaction term between the relative price of gasoline and the relative population (N^{CA}/N^{BC}). Additionally, measures of per capita income in the localities under analysis are included: I^{CA} represents California's income and I^{BC} represents Baja California's income. The population in each state (N^{BC} and N^{CA}), the number of cars in BC (C), and the number of travelers crossing the border between California (CA) and BC (V). Thus, the specification of the demand for gasoline (G) in BC is:

$$\begin{aligned} \ln G_t = & \alpha + \beta_1 \ln(P_t^{BC}) + \beta_2 \ln\left(\frac{P_t^{CA}}{P_t^{BC}}\right) + \beta_3 \left[\ln\left(\frac{P_t^{CA}}{P_t^{BC}}\right) \right] \left[\ln\left(\frac{N_t^{CA}}{N_t^{BC}}\right) \right] + \beta_4 \ln\left(\frac{I_t^{BC}}{N_t^{BC}}\right) + \beta_5 \\ & \ln\left(\frac{I_t^{CA}}{N_t^{CA}}\right) + \beta_6 \ln(N_t^{BC}) + \beta_7 \ln(N_t^{CA}) + \beta_8 \ln(V_t) + \beta_9 \ln(C_t) + e_t \end{aligned} \quad (3)$$

In equation (3), $\ln$ represents the natural logarithm, the subscript t refers to the month, and e_t represents the error term. This specification, however, presented severe multicollinearity problems that biased the estimated parameters and the hypothesis tests. A clear indication of this can be seen in the correlation matrix in Table 2. For reference purposes only, specification (3) was estimated using ordinary least squares (OLS), the constraint $\beta_1 = \beta_3 = 0$ was imposed, and the income figure was taken without normalization by the populations of BC and CA. The restriction on the parameters implies that the direct price elasticity (β_2) is equal to the negative of the cross-price elasticity of gasoline demand ($-\beta_2$).

In the context of this article, the use of time-series data requires verifying whether the variables used in equation 3 have a unit root. If this is the case, and once it has been demonstrated that there is a stable long-term correlation between the variables in question (cointegration), the following specification will be estimated:

$$\ln(G_t) = \alpha + \sum_{i=1}^p \beta_{1,i} \ln(G_{t-p}) + \sum_{i=1}^q \beta_{2,i} \ln\left(\frac{P_{t-q}^{CA}}{P_{t-q}^{BC}}\right) + \beta_3 \ln(V_t) + \beta_4 \ln(C_t) + \beta_5 \ln(N_t^{BC}) + \beta_6 \ln(N_t^{BC}) + \beta_7 \ln(I_t^{BC}) + \beta_8 \ln(I_t^{CA}) + \lambda TCE_{t-1} + e_t \quad (4)$$

The subscripts p and q denote the number of lags of the dependent variable and the variable $\ln\left(\frac{P_{t-q}^{CA}}{P_{t-q}^{BC}}\right)$ used in (4). The error correction term $\lambda = (1 - \sum_{i=1}^p \delta_i)$ is the speed of adjustment parameter, while TCE_{t-1} captures the long-run relation of the model. If $\lambda < 0$, then the model exhibits cointegration. In that case, the long-run elasticity with respect to the relative price of gasoline is calculated by dividing the parameter of the natural logarithm of that variable (β_2) by λ .

As noted in the literature review, other studies on Mexico that have estimated the effect of the relative price of gasoline on demand in the northern border region include those by Haro López and Ibarrola Pérez (1999), Ayala Gaytán and Gutiérrez González (2004) and Fullerton et al. (2012).

Estimation

The results obtained by estimating equation (3) using OLS—assuming that the variables do not exhibit unit root problems—are discussed first, thereby establishing a frame of reference that illustrates the approach followed in the demand estimation process. In these estimates, multicollinearity diagnostics were performed using the correlation matrix reported in Table 2 and the variance inflation factors (VIF). In addition, the Breusch-Pagan test was applied to detect heteroscedasticity, and the Breusch-Godfrey test was applied to detect autocorrelation in the models.

Subsequently, the augmented Dickey-Fuller (ADF) test was conducted to verify the presence of a unit root in each variable. As discussed in the results section, the variables exhibited different orders of integration, which led to the application of the cointegration method proposed by Pesaran et al. (2001), involving the estimation of autoregressive models with distributed lags.

The choice of the autoregressive distributed lag (ARDL) model to empirically represent the gasoline market in Baja California is based on both economic and technical

considerations. From an economic perspective, gasoline demand is a dynamic process: consumers do not instantly adjust their behavior in response to a price change, as such an adjustment involves modifying consumption habits, travel patterns and, in the long run, decisions regarding the vehicle fleet. This inertia in consumption justifies incorporating lags of the dependent variable into the demand equation.

Likewise, gasoline prices—both domestic prices and those on the California side of the border—can affect demand with a time lag, which justifies the inclusion of lags of the explanatory variables. The ARDL model makes it possible to capture these patterns of dynamic adjustment flexibly. That is, the use of ARDL acknowledges the natural gradual adjustment that gasoline demand undergoes in response to changes in the independent variables and distinguishes between short- and long-run effects.

From a technical standpoint, the ARDL framework is particularly appropriate because the variables in the model have different orders of integration: some are integrated of order zero ($I(0)$), and others are integrated of order one ($I(1)$). The cointegration method based on the bounds test by Pesaran et al. (2001)—designed specifically for the ARDL framework—is valid regardless of the order of integration of the regressors, unlike the classical Engle-Granger or Johansen approaches, which require that all variables be $I(1)$.

In addition, the ARDL model, in its error correction form, provides a direct way to distinguish between short- and long-run relations: the parameters associated with the first differences of the variables capture the short-run effects (dynamic adjustment), while the parameters of the error correction term (ECT) capture the long-run equilibrium relation. Long-run elasticities are obtained by dividing the coefficients of the level of the variables by the adjustment speed parameter λ , as described in equation (4).

In this way, ARDL provides robust and unbiased estimators, which are ideal for small sample sizes. The method also offers greater flexibility in the dynamic structure of the empirical model, as it allows for different numbers of lags for each variable and enables the identification of long-run relations based on short-run dynamics (Kripfganz & Schneider, 2023).

To identify the most appropriate ARDL model specification, a selection strategy known as “general-to-specific” is followed. This procedure starts with a general specification that includes the maximum number of lags for the dependent variable and the explanatory variables. It then progressively moves toward more parsimonious specifications that eliminate nonsignificant terms. In this study, the Schwarz information criterion (BIC, Bayesian information criterion) serves as a model selection tool: since the BIC penalizes overparameterization more strictly than the Akaike information criterion (AIC), it favors more parsimonious models and, therefore, models with greater generalization capacity. Multiple combinations of lags were explored for the dependent variable and for the relative price of gasoline, and the combination that minimized the BIC was selected. The results of this process are reported in the results section.

Data

The database used to estimate demand (3) is a monthly time series spanning from January 1997 to December 2015. The availability of information on gasoline demand in BC and on gasoline prices at the BC border limited the scope of the database. Additionally, as noted above, the last year in which the Mexican federal government set gasoline prices was 2015. Beginning in 2016, a system of differentiated prices was implemented in different regions of the country to prepare for the eventual liberalization of prices in 2017.

The dependent variable is the monthly volume of domestic sales of Magna gasoline, in liters, in the municipalities of Rosarito, Ensenada and Mexicali in Baja California. The source of the data is the Energy Information System of Petróleos Mexicanos (Pemex). In turn, the monthly price of Magna gasoline, in pesos per liter at the northern border, was obtained from the Energy Information Bank of the Instituto Nacional de Estadística y Geografía (Inegi).

The monthly prices per liter for BC and CA are expressed in MXN. To convert these prices from USD to MXN, the daily monthly average of the FIX MXN-USD exchange rate was used (Sistema de Información Económica, n. d.). The price for BC corresponds to the price of Magna gasoline in Mexico's northern border region, while the price for CA corresponds to the average price of regular gasoline throughout the state of California.

Per capita income in Baja California was calculated as the ratio of the state's annual gross domestic product (GDP)—in millions of 2013 constant Mexican pesos—to Baja California's annual population. Per capita income in California, was calculated as the ratio of the state's annual GDP—in millions of 2012 Mexican pesos—to California's annual population. The per capita GDP data for both states are annual figures. To construct the database used in the estimation of demand, the annual figure was repeated for each month of the corresponding year.

The variable for travelers crossing the border refers to the number of private cars, buses and cargo vehicles entering the United States from Mexico at the Andrade, Calexico, Calexico East, Otay Mesa, San Ysidro and Tecate border crossings.

The vehicle fleet was calculated by adding the number of cars, buses, trucks and motorcycles in circulation in the municipalities of Ensenada, Mexicali, Tecate and Tijuana in Baja California for each year.

Table 1 describes the variables used in this article, as well as the sources of information, while Table 2 presents the descriptive statistics for the variables.

Table 2. Descriptive statistics and correlation matrix

Variable	G_t	P_t^{BC}	P_t^{CA}	N_t^{BC}	N_t^{CA}	I_t^{BC}	I_t^{CA}	V_t	C_t
Minimum	68853840	2.790	2.458	2256820	32486010	0.1303	0.2565	1012172	449521
Maximum	183465570	13.570	15.812	3357794	38918045	0.1825	1.0847	3249254	1085329
Mean	138771912	7.227	8.043	2887888	35973021	0.1517	0.5567	2564065	775526
Sta. Dev.	25299724	2.801	3.961	348704.4	1868019	0.0128	0.1893	327777.2	180305.7
G_t	1.00								
P_t^{BC}	0.66	1.00							
P_t^{CA}	0.83	0.94	1.00						
N_t^{BC}	0.78	0.94	0.95	1.00					
N_t^{CA}	0.72	0.95	0.94	0.99	1.00				
I_t^{BC}	-0.72	-0.61	-0.70	-0.77	-0.74	1.00			
I_t^{CA}	0.68	0.94	0.91	0.94	0.96	-0.64	1.00		
V_t	-0.43	-0.30	-0.37	-0.39	-0.27	0.38	-0.25	1.00	
C_t	0.43	0.45	0.49	0.62	0.58	-0.54	0.52	0.14	1.00

Note. All Pearson correlation coefficients are significant (p – value < 0.05).

Source: created by the authors

Results

Table 3 presents the OLS estimates of the models and groups the independent variables into blocks to assess the consistency of the results. The last column of the table shows the VIF for each variable. Models 1, 2 and 5 do not exhibit multicollinearity issues. As shown in Table 3, the variables associated with the population are responsible for the multicollinearity. The Shapiro-Wilks test, in which no model rejected its null hypothesis, leads to the conclusion that the residuals of each model follow a normal distribution. Nonetheless, the Breusch-Pagan and Breusch-Godfrey tests found that all models exhibit heteroscedasticity and autocorrelation; therefore, the Eicker-Huber-White standard error correction was applied to the estimates.

Table 3. Results of ordinary least squares model estimates

Variable	Model 1	Model 2	Model 3	Model 4	Model 5	vif
Constant	18.699*** (0.006)	21.733*** (1.283)	98.370*** (9.436)	113.854*** (11.047)	21.383*** (0.796)	
$\ln\left(\frac{P_t^{BC}}{P_t^{CA}}\right)$	-0.761*** (0.029)	-0.719*** (0.042)	-0.419*** (0.037)	-0.413*** (0.037)	-0.587*** (0.037)	[2] 1.607 [3] 2.955 [4] 2.971 [5] 2.056
$\ln(V)$		-0.219** (0.093)	-0.054* (0.029)	-0.057* (0.031)	-0.097* (0.056)	[2] 1.182 [3] 1.356 [4] 1.530 [5] 1.463
$\ln(C)$		0.015 (0.029)	-0.265*** (0.027)	-0.264*** (0.027)	-0.133*** (0.032)	[2] 1.557 [3] 3.018 [4] 3.037 [5] 2.279
$\ln(N_t^{BC})$			4.349*** (0.415)	4.092*** (0.424)		[3] 156.94 [4] 171.38
$\ln(N_t^{CA})$			-8.045*** (0.882)	-8.723*** (0.927)		[3] 125.80 [4] 147.11
$\ln(I_t^{BC})$				-0.143* (0.084)	-0.344*** (0.093)	[4] 3.266 [5] 2.730
$\ln(I_t^{CA})$				0.173*** (0.061)	0.145*** (0.020)	[4] 27.167 [5] 2.620
AIC	-449.764	-472.185	-608.437	-613.643	-537.457	
BIC	-439.476	-455.039	-584.432	-582.778	-513.452	
R ² Aj	0.770	0.793	0.887	0.891	0.846	
Est F	760.153***	291.171***	358.037***	265.173***	250.378***	
χ^2	113.07***	77.758***	57.174***	53.143***	79.617***	
W	0.992	0.992	0.994	0.992	0.996	

Notes: 228 observations. vif is the variance inflation factor. Eicker-White standard errors are shown in parentheses. *Significance level of 0.1; **0.05; ***0.01. AIC is the Akaike information criterion; BIC is the Schwartz information criterion. The F-statistic is for the joint significance test. χ^2 corresponds to the Breusch-Godfrey test statistic for autocorrelation. W is the Shapiro-Wilks test statistic for the normality of the errors.

Source: created by the authors

The results obtained by applying OLS to models 1, 2 and 5 indicate that the parameter for the relative price of Mexican gasoline is statistically significant. The coefficients of determination (R^2) range from 0.77 to 0.846. In model 1, the price elasticity of gasoline demand at the Baja California border is -0.761 ; that is, a 10% increase in price would reduce demand by 7.61%. Furthermore, according to the estimated specification for model 1, the price elasticity of demand in Baja California with respect to the price of gasoline in California is 0.761.

The OLS estimates also show that the elasticity of demand for BC with respect to California (I_t^{CA}) income is statistically different from zero, with an estimated value of 0.145 in model 5. On the other hand, unexpected results were obtained regarding Baja California income and the number of vehicles crossing the border. Similarly, although the number of vehicles in circulation in Baja California shows a direct correlation in model 2, its parameter is not statistically significant. These results may be attributable to the fact that the independent variables used to estimate income and the number of registered vehicles correspond to the entire state of Baja California and not specifically to its border region.

Table 4 shows the results of the ADF test and the integration order of each variable. Variables that did not exhibit a unit root in their logarithmic form were identified as integrated of order zero. The remaining variables were transformed into first differences, and all were found to lack a unit root; they were identified as integrated of order one.

Estimating the baseline empirical model requires that the variables be cointegrated. Given the presence of multiple orders of integration, in accordance with the method of Pesaran et al. (2001), the autoregressive distributed lag (ARDL) model with error correction was used. In particular, the variables for revenue (I_t^{BC} and I_t^{CA}), population (N_t^{BC} and N_t^{CA}), vehicle fleet (C_t), and the number of vehicles crossing the international bridges in Baja California (V_t) were considered exogenous to determine the order of lags in which they should appear in the specification, which implies that they were included only in their contemporary order. This is due to the way these variables were constructed.

When applying Schwarz's information criterion (BIC), the models used thirteen lags of the dependent variable and no lags of $\ln\left(\frac{P_t^{BC}}{P_t^{CA}}\right)$. With this specification and in accordance with Pesaran et al. (2001), the Wald boundary test for the absence of cointegration was conducted for each estimated model. The results, together with the estimated parameter for the error correction term—which is less than zero—confirm the presence of cointegration in all models. This makes it possible to estimate the parameters and calculate the long-run elasticities. The results are presented in Table 5.

Table 4. Augmented Dickey-Fuller tests for unit roots

ADF test	Intersection and trend			Trend only		None
Variable	τ_3	φ_3	φ_2	τ_2	φ_1	τ_1
$\ln(G_t)$	-2.578	2.298	3.335	-1.752	1.644	0.450
$\Delta \ln(G_t)$	-16.518***	90.975***	136.462***			
$\ln\left(\frac{P_t^{BC}}{P_t^{CA}}\right)$	-3.913***	5.120**	7.679**			
$\ln(V_t)$	-9.485***	31.559***	46.577***			
$\ln(C_t)$	-1.615	1.219	1.534	-1.701	1.743	0.7346
$\Delta \ln(C_t)$	-10.587***	37.361***	56.941***			
$\ln(N_t^{BC})$	-1.549	7.789***	2.502	-1.863	10.895***	4.239
$\Delta \ln(N_t^{BC})$	-12.044***	48.356***	72.533***			
$\ln(N_t^{CA})$	-3.145	10.269***	5.4026	-1.326	10.534***	4.382
$\Delta \ln(N_t^{CA})$	-11.985***	47.887***	71.830***			
$\ln(I_t^{BC})$	-1.568	0.962	1.406	-1.490	1.147	0.207
$\Delta \ln(I_t^{BC})$	-10.542***	37.047***	55.571***			
$\ln(I_t^{CA})$	-3.305	6.659**	5.462	-0.652	4.554*	-2.901***

Note: *Significance level of 0.1; **0.05; ***0.01

Source: created by the authors

It is worth noting that the optimal specification identified by the BIC corresponds to an ARDL (13, 0) model—that is, thirteen lags of the dependent variable and zero additional lags of the relative price of gasoline in the short-run dynamics—. Far from being a problematic result, this finding is economically and statistically consistent. First, the data are monthly, and gasoline demand exhibits marked seasonal dynamics, so that the lags of the dependent variable capture the cyclical patterns of consumption throughout the year. Second, the fact that the relative price does not require lags in the short-run component does not imply that it has no effect: the relative price does appear in the level of the error correction term (ECT), where its coefficient captures the long-run relation. The “general-to-specific” strategy implemented using the BIC, which penalizes the inclusion of redundant parameters more severely than the AIC, leads to the elimination of price lags in short-run dynamics because their marginal contribution does not offset the cost of increased parameterization. The model remains an ARDL model with error correction, and the long-run elasticities are correctly estimated based on the ratio of the relative price coefficient in the ETC to the adjustment speed parameter λ .

Table 5. Short-run results of the ARDL estimates

Variable	Model 6	Model 7	Model 8	Model 9	Model 10
Constant	4.616*** (0.856)	3.130** (1.286)	14.851 (9.602)	13.757 (12.088)	2.523** (1.286)
Short-run estimate					
$\Delta \ln(G_{t1})$	-0.419*** (0.065)	-0.463*** (0.064)	-0.427*** (0.068)	-0.446*** (0.074)	-0.481*** (0.068)
$\Delta \ln(G_{t2})$	-0.385*** (0.062)	-0.427*** (0.064)	-0.397*** (0.066)	-0.415*** (0.069)	-0.443*** (0.067)
$\Delta \ln(G_{t3})$	-0.228*** (0.063)	-0.265*** (0.061)	-0.240*** (0.061)	-0.256*** (0.064)	-0.277*** (0.063)
$\Delta \ln(G_{t4})$	-0.321*** (0.054)	-0.353*** (0.055)	-0.334*** (0.056)	-0.348*** (0.056)	-0.364*** (0.056)
$\Delta \ln(G_{t5})$	-0.128** (0.053)	-0.162*** (0.054)	-0.148*** (0.055)	-0.162*** (0.057)	-0.175*** (0.056)
$\Delta \ln(G_{t6})$	-0.224*** (0.054)	-0.248*** (0.055)	-0.239*** (0.057)	-0.252*** (0.056)	-0.257*** (0.054)
$\Delta \ln(G_{t7})$	-0.092* (0.049)	-0.107** (0.049)	-0.105** (0.052)	-0.166** (0.051)	-0.144** (0.050)
$\Delta \ln(G_{t8})$	-0.122*** (0.045)	-0.147*** (0.048)	-0.149*** (0.052)	-0.163*** (0.049)	-0.162*** (0.048)
$\Delta \ln(G_{t9})$	-0.031 (0.046)	-0.046 (0.047)	-0.051 (0.049)	-0.061 (0.049)	-0.058 (0.049)
$\Delta \ln(G_{t10})$	-0.281*** (0.068)	-0.293*** (0.068)	-0.298*** (0.069)	-0.306*** (0.069)	-0.304*** (0.068)
$\Delta \ln(G_{t11})$	-0.261*** (0.091)	-0.254*** (0.091)	-0.260*** (0.093)	-0.263*** (0.091)	-0.259*** (0.089)
$\Delta \ln(G_{t12})$	0.257*** (0.087)	0.242*** (0.086)	0.238*** (0.084)	0.227*** (0.085)	0.227*** (0.087)
Error correction term					
$\ln(G_{t-1})$	-0.247*** (0.046)	-0.209*** (0.054)	-0.254*** (0.081)	-0.244*** (0.084)	-0.203*** (0.056)

Tabla 5. Continued

Long-run estimate				
$\ln \left(\frac{P_t^{BC}}{P_t^{CA}} \right)$	-0.247*** (0.040)	-0.241*** (0.041)	-0.234*** (0.038)	-0.234*** (0.038)
$\ln(V_t)$		0.063* (0.033)	0.061* (0.034)	0.079** (0.038)
$\ln(C_t)$		-0.011 (0.012)	-0.032 (0.023)	-0.033 (0.024)
$\ln(N_t^{BC})$			0.540 (0.441)	0.374 (0.469)
$\ln(N_t^{CA})$			-1.069 (0.838)	-0.899 (0.987)
$\ln(I_t^{BC})$				-0.083* (0.050)
$\ln(I_t^{CA})$				0.016 (0.034)
Short-run elasticity				
$\mathcal{E} \frac{P_t^{BC}}{P_t^{CA}}$	-0.2466	-0.2409	-0.2340	-0.2344
Long-run elasticity				
$\mathcal{E} \frac{P_t^{BC}}{P_t^{CA}}$	-0.9990	-1.1522	-0.9209	-0.9612
AIC	-8.24.8413	-824.8363	-823.1630	-821.9609
BIC	-770.9111	-764.1648	-755.7502	-747.8069
Adjusted R^2	0.751	0.753	0.753	0.754
F-statistic	47.137***	41.831***	37.326***	33.813***
χ^2	25.277	26.456	23.972	23.275
F_W	20.971***	23.254***	20.301***	20.381***

Notes: 215 observations. Eicker-White standard errors in parentheses. *Significance 0.1; **0.05; ***0.01. AIC is the Akaike information criterion; BIC is the Bayesian information criterion. χ^2 corresponds to the statistic from the Jung-Box test for normality of errors. F_W corresponds to the statistic from the Wald bounds test for the absence of cointegration.

Source: created by the authors

The Breusch-Pagan test was also performed on these models, and the null hypothesis of no heteroscedasticity was rejected in all cases; therefore, consistent estimators of the errors (Eicker-White standard errors) were used. In addition, the Ljung-Box test was conducted to verify the absence of autocorrelation in the residuals; since the null hypothesis was not rejected, it is concluded that these models do not exhibit autocorrelation.

In addition, the results presented provide evidence of the stability of the estimated parameters. First, the remarkable consistency of the short-run elasticity across the five models in Table 5—whose value ranges from -0.234 to -0.247 —indicates that the estimates are not sensitive to the inclusion or exclusion of additional control variables, which constitutes evidence of structural robustness. Second, the adjustment speed parameter λ is negative and statistically significant in all models, confirming that the system converges to long-run equilibrium and ruling out the possibility of dynamic instability. Third, the Wald boundary test (F_w statistic in Table 5), which is significant in all models, confirms the existence of a stable cointegration between the variables. Taken together, these results provide a solid basis to claim the stability of the model and the robustness of the long-run conclusions.²

In the five models estimated using ARDL, the parameter for the natural logarithm of the relative price of gasoline is statistically significant and has the expected sign. It is interesting to note the consistency of the estimated parameter across the models in Table 5, where its value ranges from -0.234 to -0.247 . Using model 6 as a reference, the short-run direct price elasticity of gasoline in Baja California is -0.247 , while the cross-price elasticity is 0.247 . This means that a 10% reduction in the price of gasoline in California would lead to a 2.47% decrease in gasoline volume in Baja California. This would be the estimated cross-border effect on the short-run demand for gasoline in Baja California.

Long-run price elasticities incorporate the effect of a one-period lag in demand adjustment. This time lag makes gasoline demand less inelastic compared to the short run, which amplifies the cross-border effect. As shown at the bottom of Table 5, the estimated relative price elasticity of gasoline demand ranges from 0.92 to 1.17 in absolute terms. Regarding the effect of gasoline prices in California—which is the main topic of this article—the estimate in model 6 indicates that a 10% reduction in gasoline prices in California would reduce gasoline volume in Baja California by 10%. This long-run estimate represents what is known as unitary demand in terms of elasticity. Thus, in the long run, the border effect is nearly five times greater than in the short run.

These results, obtained using time-series techniques applied to a model modified from the specification by Banfi et al. (2005), contrast with the findings of Ibarra Salazar and Sotres Cervantes (2008, 2022), who estimated a somewhat different specification of the demand models using a panel dataset covering all states of Mexico. On the other hand, studies using monthly time series, such as that by Fullerton et al. (2012), that estimated a short-run relative price elasticity for Ciudad Juárez, Chihuahua, of 0.26. Ayala Gaytán and Gutiérrez González (2004) reported a similar elasticity of 0.131

² An alternative method used for time series analysis is dynamic ordinary least squares (DOLS). In this case, it was considered unnecessary to apply this method. The ARDL-ECM estimator proposed by Pesaran et al. (2001) produces long-run parameters that are asymptotically efficient and consistent under cointegration, since the error correction term simultaneously accounts for the endogeneity of the regressors and for serial correlation in the residuals—properties it shares with DOLS. The comparative advantage of DOLS over ARDL is most evident when all variables are $I(1)$ and the sample size is small; however, in this case, the variables have mixed integration orders of $I(0)/I(1)$, a condition under which the ARDL model proposed by Pesaran et al. (2001) is the preferred estimator in the literature. Furthermore, with 228 monthly observations, the sample size is sufficient for the asymptotic properties of the ARDL-ECM to provide a good approximation. For these reasons, internal consistency of the results in Table 5—in terms of signs, magnitudes and statistical significance across the five models—together with the evidence of cointegration and stability described above, is deemed sufficient to support the long-run conclusions without requiring an additional DOLS estimation.

for Tijuana, Rosarito and Tecate, BC (Zone IA), and 0.17 for Mexicali, BC (Zone IB), with long-run elasticities estimated at 0.13 and 0.17, respectively. In addition, Haro López and Ibarrola Pérez (1999) found a short-run elasticity of 0.25 for Zone IA and 0.15 for Zone IB.

The parameter for Baja California's income, included in models 9 and 10, is significant in both cases, although it does not have the expected sign. Models 4 and 5 also show an estimated income elasticity for Baja California that is negative and statistically different from zero. In contrast to the ordinary least squares estimate, the parameter for the natural logarithm of California's income was not statistically significant. An inverse correlation between income and gasoline demand is considered counterintuitive. As a result, in some cases, such findings are not reported in empirical studies (Havranek & Kokes, 2015). Dahl's (2012) analysis includes 1 091 studies that estimated income elasticities for gasoline. Among them, about 10% report negative income elasticities. Of these, 80% involve models that include the number of cars as an explanatory variable. The inverse correlation between income and gasoline demand suggests that as income rises, consumers can afford more fuel-efficient modes of transportation.

Regarding the number of vehicles crossing the border (V), the results are now consistent, indicating that the coefficient for this variable is positive and statistically significant. The estimated elasticity of gasoline demand at the Baja California border with respect to V ranges from 0.061 to 0.092.

Conclusions

This paper contributes to the literature on cross-border shopping in goods when regulation, market structure or tax policy differs on either side of a geographic border between jurisdictions within the same country or between countries. In the case of the Mexico-United States border, gasoline tourism is a constant phenomenon, with periods when the price differential for gasoline on both sides of the border is so large that it motivates consumers to travel to the other country to fill up their tanks.

Demand in the Baja California border region was estimated using a monthly time series from January 1997 to December 2015. By applying the distributed lag autoregressive method, and based on the demand specification in Banfi et al. (2005), the elasticities of gasoline demand in the Baja California border region were estimated with respect to the price of gasoline in that region (direct price elasticity) and with respect to the price of gasoline in California. It was found that, in the short run, demand for gasoline at the Baja California border is inelastic with respect to both prices. Specifically, evidence was found that if the price of gasoline in California were to fall, for example, by 10%, demand for gasoline at the Baja California border would decrease by 2.5% in the short run and by 10% in the long run. The evidence shows that this effect is statistically significant.

The estimation of the gasoline border effect described in this article provides evidence that, at the time, the policy of harmonizing gasoline prices and the fiscal policy (consisting of the weekly adjustment of the IEPS [special tax on production and services] and the application of a reduced VAT [value-added tax] rate of 8%) toward the gasoline sector in the border region made sense. Small variations in gasoline prices on the northern side of the border would result in significant reductions in gasoline demand in border cities.

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