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Short-run price and income elasticity of gasoline demand: Evidence from Lebanon

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ABSTRACT

We empirically estimate the demand for gasoline in the presence of multiple shifts caused by structural breaks using monthly data from Lebanon covering the period 2000:M1–2010:M12. Consistent with most studies in the literature, our study reports that gasoline is price and income inelastic in the short-run. However, when a single and multiple breaks are introduced, the consumers' responsiveness to gasoline price and income increase. Since both price and income elasticity are sensitive to structural changes, a policy that pleads for a flat excise tax may not be optimal with respect to either the cyclical pattern of government revenues or the internalization of international environment standards.

1. Introduction

A large number of studies have investigated the relationship between gasoline consumption, price and income. While these studies differ in scope and method, price and income elasticity of gasoline demand tend to be inelastic both in the short- and the long-run regardless of whether a country is a major oil producer (e.g., Al-Sahlawi, 1988; Eltony, 1994; Eltony and Al-Mutairi, 1995; Eltony, 1996; and Crôtte et al., 2010) or a major oil importer (e.g., Graham and Glaister, 2002; and Hughes et al., 2008).

Dahl and Sterner (1991) surveyed the literature and provided an average short-run price elasticity of gasoline demand of -0.26 and an average short-run income elasticity of gasoline demand of 0.48, while Espey (1996) in his encompassing survey over price elasticities in the US and other industrialized countries provides a median of -0.23 and 0.39 for short-run price and income elasticity, respectively. Contrasting these elasticities with those of some of the GCC countries, the price and income responsiveness of Kuwaitis (Eltony and Al-Mutairi, 1995) and Saudis (Al-Sahlawi, 1988) to variation in gasoline demand are -0.37 and 0.47, and -0.08 and 0.11, respectively. While for Saudi Arabia elasticities are consistent with its industrial structure and government policies, elasticities in Kuwait are surprisingly in line with elasticities in industrialized countries, which are in general net importers of oil. Hence, the low responsiveness of consumers to variation in gasoline price and disposable income may depend on technological, economic, and political structural changes (see for example, Espey, 1998; Sterner, 2007; and Hughes et al., 2008).

The literature on gasoline demand can be broadly subdivided into two main strands in terms of the estimation approach. Because consumption, price and income are typically non-stationary variables, one approach uses cointegration for non-stationary variables and looks at the long-run and short-run relationship between consumption and price within the Error Correction Model (ECM) (e.g., Bentzen, 1994; Ramanathan, 1999; Eltony and Al-Mutairi, 1995; and Badr et al., 2008). Another approach is to use the Partial Adjustment Model (PAM) for de-trended variables (e.g., Hughes et al., 2008; and Wadud et al., 2010).

As a matter of fact, different situations may arise with a closer examination of the statistical properties of the variables used. First, when all the variables are non-stationary, a combination of cointegration and ECM can be used to estimate elasticities of gasoline demand both in the short- and long-run. Second, when all the variables are stationary, the PAM can be used to estimate short-run elasticities on the basis of which long-run elasticities are also computed. Finally, when at least one of the variables is stationary while the rest are not, both ECM and PAM estimations will become questionable.

Perron (2006) gives a full measure of the three cases in terms of the identification of nonstationary and stationary processes, and the modeling of these processes. Taking structural breaks into consideration when testing for unit root and stationarity implies that the test is information-based as the length of the time trend (under the null hypothesis of no structural break) is shortened

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to disentangle the effect of one structural break from another. We benefit from his insight in this paper to study the case of Lebanon, a country that has experienced a number of dramatic political events in its recent history.

Against this background, the purpose of this study is to investigate short-run price and income elasticity when multiple endogenous breaks are considered. Few studies have integrated the notion that elasticities can be changing structurally (e.g., Espey, 1998; and Hughes et al., 2008). We use both a deterministic and a stochastic approach to capture the effects of structural breaks on nonstationary processes, and short-run price and income elasticities.

Previous studies have used structural breaks in identifying whether a nonstationary process under the null hypothesis of no structural break (under the standard unit root tests) is still nonstationary when structural breaks are considered (e.g., Lee and Chang, 2005; and Kumar and Shahbaz, 2010). However, these studies do not investigate the influence of structural changes on price and income elasticities. Our approach is related to Hughes et al. (2008) who consider two different periods to investigate whether elasticities are constant or structurally changing as suggested in Espey (1998). However, we differ from Hughes et al. (2008) in that our structural breaks are endogenous, which is in line with the literature on the identification of the dates when a structural change is suspected to have occurred (e.g., Bai and Perron, 2003; and Perron, 2006).

Our findings are in line with Hughes et al. (2008) in that elasticities are changing with structural breaks. However, in our case, elasticities tend to be higher the more the number of breaks is included in a typical gasoline demand model. These high elasticities cast forwards a strong contrast on elasticities of a model ignoring structural breaks and a model including only a single structural break. Because elasticity estimates vary according to the empirical model specification, we report here (for comparison reasons) the elasticity on a model ignoring structural breaks and a model including a single structural break. The price and income elasticity are -0.623 and 0.309 in a model excluding breaks, respectively; while they are -0.915 and 0.424 in a model with a single structural break, respectively. We conclude based on these figures that price and income elasticities are structurally changing and that structural breaks enhance the responsiveness of consumers to changes in price and income. A direct implication of our findings is that policy makers can better time their decisions on gasoline taxation and its impact on income distribution and government revenue. Indeed, policy making can be improved by using a model that includes either a single structural break when the data span is shorter or multiple breaks when the data span is longer with respect to changes in technology, industrialization, economic integration or democratic practices.

The rest of the paper unfolds as follows. The next section describes the data and discusses the econometric methods and the empirical results. Finally, concluding remarks are reported in Section 3.

2. Data and model specifications

In this section, the data set and the empirical specification and results are presented.

2.1. Data

We use monthly data on gasoline consumption (Q_t), total imports in US dollars (M_t), population (N_t), consumer price index (Θ_t), private spending in US dollars (Y_t), and nominal gasoline prices in US dollars (P_t).[1] The data sample runs from January 2000 to December 2010 for a total of 132 observations. We obtain M_t and Y_t from the Central Bank of Lebanon,[2] P_t from the Ministry of Energy and Water (Lebanon),[3] Θ_t from the Consultation and Research Institute (Beirut)[4] and Q_t and N_t from the Central Administration of Statistics (Lebanon).[5]

Table 1
Descriptive statistics.

	q_t	p_t	m_t	y_t
Average	0.597	1.913	4.558	3.072
Standard deviation	0.762	0.184	0.315	0.668
Maximum	3.194	2.256	5.252	4.088
Minimum	−1.717	1.494	3.442	1.556
Skewness	0.956	−0.320	−0.084	−0.750
Excess kurtosis	3.085	−0.162	0.036	−0.595
Jarque–Bera	72.436*	2.399	0.162	14.329*
Autocorrelations and Ljung–Box statistics				
Lag 1	0.457	0.967	0.814	0.969
Lag 2	0.447	0.932	0.763	0.965
Lag 4	0.384	0.878	0.739	0.962
Lag 10	−0.128	0.803	0.693	0.918
Lag 12	−0.048	0.798	0.700	0.902
Ljung–Box statistic at lag 12	112.416*	1228.676*	857.638*	1514.195*

q_t is gasoline consumption per capita, p_t is real price per 20 l, m_t is real income in terms of imports, y_t is income in terms of private spending, and * means that the statistic is significant at 5%.

Given that figures on disposable income are not available at any frequency in Lebanon, we use both monthly data on import and private spending as proxy for disposable income. These two proxies capture monthly variations in income based on the following motivations. First, private spending is computed as purchase inside Lebanon by residents and nonresidents (I_{1t}) plus ATM cash withdrawals inside Lebanon by residents and nonresidents (I_{2t}) minus purchases and ATM cash withdrawal outside Lebanon by residents (I_{3t}) or $Y_t = I_{1t} + I_{2t} - I_{3t}$. Second, the use of import as a proxy for the Lebanese national income is justified given that the correlation between GDP and imports over the sample period was found to be 86%. Furthermore, Persen (1958) reported that the share of imports to GDP was 47%, and 54% in 1951 and 1956, respectively. This trend has been maintained since as this share was 48% in 2007. Additionally, imports have been used as proxy for GDP in some of studies on the Lebanese economy (e.g., Nasr et al., 2000).

Let q_t be equal to $\ln(Q_t/N_t)$, $p_t = \ln(P_t/\Theta_t)$, $y_t = \ln[(Y_t/\Theta_t)N_t^{-1}]$, and $m_t = \ln[(M_t/\Theta_t)N_t^{-1}]$. It is worth noting that q_t is given in units of 20 l.[6] Table 1 reports the summary statistics on the central moments, the autocorrelation coefficients, and Ljung–Box statistics of the logarithmic variables.

Looking at the skewness and kurtosis, Table 1 shows that q_t and p_t are not normally distributed, while m_t and y_t do not deviate from normality. Nonetheless, the leptokurtic tendency is not critical when dealing with macroeconomic variables, which are aggregated at much lower frequency than for example financial variables. For such data, the first two moments are critical, as they determine whether the series carry deterministic trends affecting their long-run average and variance.

[1] During our sample period 2000:M1–2010:M12, the exchange rate of the Lebanese Pound (LBP) has been fixed against the U.S. Dollar (USD). We convert gasoline into dollar at the official peg of 1,507.50 LBP/USD.

[2] http://www.bdl.gov.lb.

[3] http://www.energyandwater.gov.lb/.

[4] http://crilebanon.com/.

[5] http://www.cas.gov.lb.

[6] This measure is the standard unit of accounting used to price gasoline in Lebanon. We convert the data on gasoline given in thousands of metric tons into unit of 20 l.

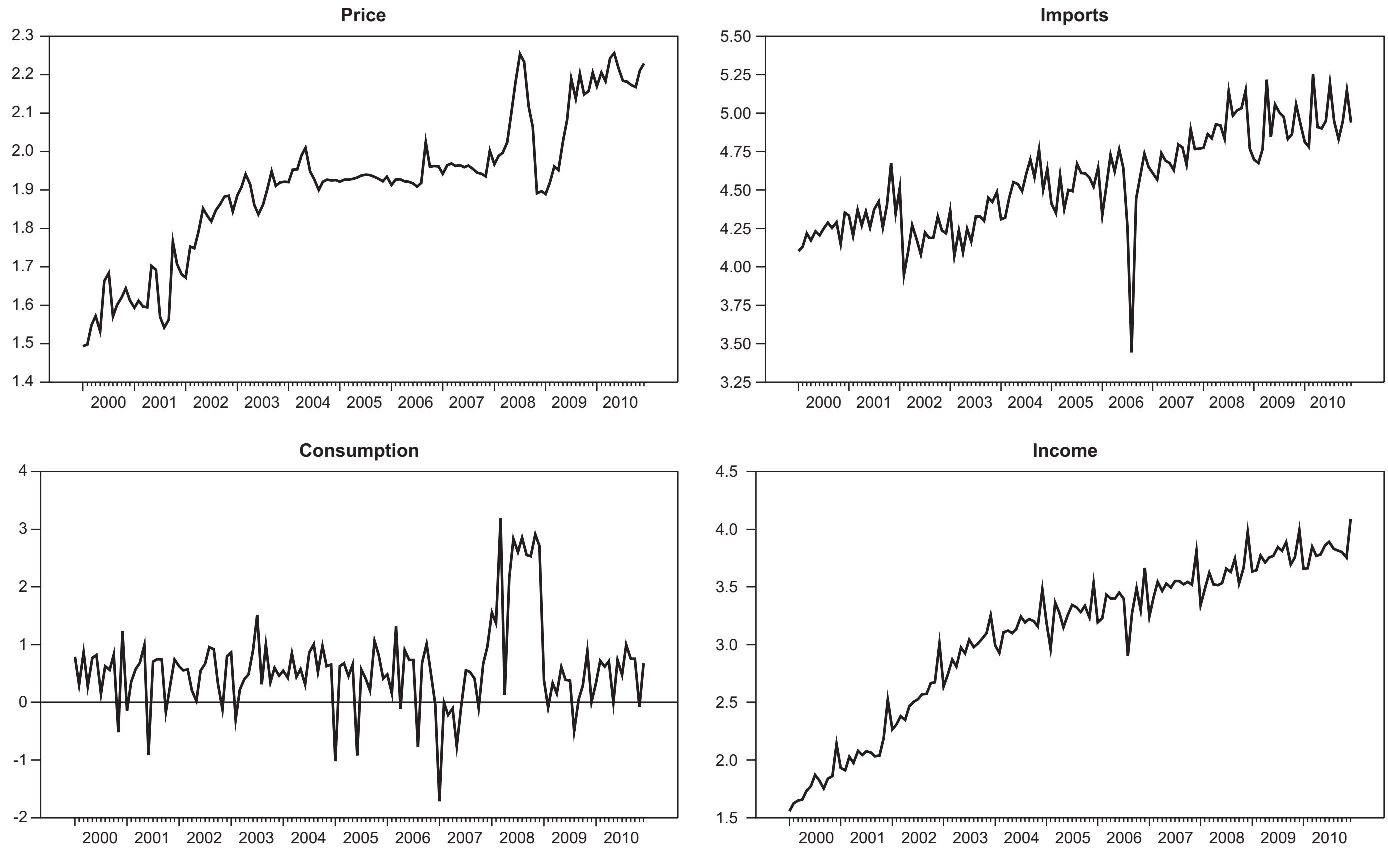

Fig. 1. Time patterns of consumption, price, income and imports.

A first glimpse at the dynamics of the long-run mean and variance is given by the 12 first autocorrelation coefficients and the Ljung–Box statistic at lag 12 of the variables. It takes much longer time for gasoline price, imports and private spending series to revert to their long-run averages when we look at the pattern of their autocorrelation coefficients and their Ljung–Box statistics at lag 12 under the null hypothesis that the 12 first autocorrelation coefficients are jointly uncorrelated. It follows from the autocorrelation functions of these variables that p_t, y_t and m_t carry over a time trend that can be both deterministic and stochastic. In contrast, the autocorrelation function of the gasoline consumption suggests a quick mean reversion, which is an indication that q_t is stationary about its long-run average. Fig. 1 gives a visual perspective of the underlying drivers of the autocorrelation functions of the four variables.

The gasoline price, import and income series exhibit typical macroeconomic properties. They show a positive trend and adjustment dynamics to normal and abnormal shocks. For example, there was a dramatic drop in import in July 2006 resulting from the inability of Lebanon to imports during the 2006-armed conflict with Israel. In contrast, upwards and downwards movements are more frequent in gasoline consumption. Nonetheless, even though both Table 1 and Fig. 1 suggest that the time series at hand deviate from their long-run averages, they don't tell us to which extent these averages increase significantly over time. Therefore, we test for the presence of a unit root in the series using the augmented Dickey–Fuller (ADF) test.

2.2. Unit-root test

Let x_t denote any of the four variables, the ADF unit-root test is based on the following regression:

$$x_t = c_t + \beta x_{t-1} + \sum_{k=1}^{p-1} \phi_k \Delta x_{t-k} + e_t, \tag{1}$$

In Eq. (1), c_t is a deterministic function of the time index t. In practice, c_t can be a constant or equal to zero. Furthermore, c_t can be given by $c_t = \omega_0 + \omega_1 t$ or it can take any other form like the one including indicators for structural breaks. Finally, $\Delta x_t = x_t - x_{t-1}$ and e_t is a white noise error term.

The null hypothesis for a unit root is $H_0 : \hat{\beta} = 1$ and the ADF statistic is given by $(\hat{\beta} - 1)/\hat{\sigma}_\beta$, where $\hat{\sigma}_\beta$ is the standard error of $\hat{\beta}$. In order for the model to be well-specified, Eq. (1) may include a deterministic component in c_t and an autoregressive component in terms of the first difference of Δx_t associated with coefficients $\hat{\phi}_k$ for $k=1$ to $p-1$ lags. Table 2 reports the ADF test.

We estimate Eq. (1) with c_t including a time trend. Moreover, we use an Akaike Information Criterion (AIC) to select the number of optimal lags in Eq. (1). The results of Table 2 are in line with the autocorrelation functions of Table 1 and the exhibits of Fig. 1. Contrasting the t-statistics on $\hat{\beta}$ with the critical value at 5% on $\hat{\beta}$, Table 2 shows that gasoline price, import and income are nonstationary processes, while gasoline consumption is a stationary process.

However, Eq. (1) does not capture features that dismantle the information structure of x_t when structural shifts (breaks) are present. There is a rich literature on whether unit root tests give definite answers on the intrinsic nature of a nonstationary process or structural breaks do alter a process in such a way that nonstationariy is less of a problem in models that combine nonstationary and stationary processes. Bai and Perron (1998) and Perron (2006) discuss the intricate interplay between nonstationary processes and processes with structural changes, which are different from processes with jumps. Structural changes tend to put a process on a higher or a lower path than before the break in such a way that the process rarely returns to its level prior to the structural break.

We examine Eq. (1) when deterministic breaks are included, and these breaks are determined endogenously. Alternatively, the

Table 2
Unit-root test (ADF).

	Optimal lag length	*t*-statistic	Critical value at 5%
log of gasoline consumption, q_t	7	−3.559*	−3.446
log of gasoline price, p_t	11	−2.162	−3.446
log of imports, m_t	12	−2.616	−3.446
log of income, y_t	12	−1.581	−3.446

The model estimated is $x_t = c_t = \beta x_{t-1} + \sum_{k=1}^{p-1} \phi_k \Delta x_{t-k} + e_t$,where x_t is any of the log variable, and one asterisk (*) means that the statistic is significant at 5%.

Table 3
Unit root test on single break and dates of breaks.

	q_t	p_t	m_t	y_t
$\hat{\omega}_0$	0.817*	0.329*	3.138*	0.752*
$\hat{\omega}_1$	9.481*	0.027	0.003	0.674*
$\hat{\omega}_2$	−0.652	−0.145*	−1.145*	−0.134
$\hat{\omega}_3$	−0.002	0.001	0.004*	0.022*
$\hat{\omega}_4$	−0.073	0.000	0.001	−0.017*
$\hat{\beta}$	−0.480*	0.795*	0.247*	0.475*
The *t*-statistic	−5.863*	−4.044	−10.825*	−4.356
Critical value at 5%	−5.550	−5.550	−5.550	−5.550
Single break dates	2008:04	2001:06	2006:07	2002:08
Number of multiple breaks	3 on	5 on	4 on	5 on
	2006:07	2001:09	2002:01	2001:11
	2007:10	2002:12	2003:09	2003:02
	2009:01	2006:08	2006:10	2004:06
		2008:03	2008:01	2006:11
		2009:06		2008:11

q_t is gasoline consumption, p_t is gasoline price, m_t is import, y_t is private spending, and * means significance at 5% or less.

breaks could be exogenously determined. The date of a break can be determined by mechanically associating it with the date of a specific socioeconomic event making headlines and reshuffling the prevalent economic or political establishment. Yet, such an approach could be misleading because these shocks may be short-lived; not permanently affecting a process in its rise or fall. For example, Fig. 1 shows that the 2006-war with Israel did impede Lebanon to import from and to export abroad, but right after the war ended in September 2006 the process reverted and continued its way upwards as if nothing happened two months earlier.

A number of procedures have been introduced to determine the dates of the break (see Perron, 2006 for a survey). We follow Bai and Perron (1998) to determine the number of breaks, the dates of the breaks and the time over which the structural effects can be seen as prevalent. As before, we denote by x_t any of the four variables at hand, and we search for a single break over the sample period. Having identified the dates of these breaks, we express c_t in terms of the breaks and a time trend as,

$$c_t = \omega_0 + \omega_1 d_t^b + \omega_2 d_t + \omega_3 t + \omega_4 d_t^a \tag{2}$$

where t is the time trend index, $d_t^b = 1$ if $t < \bar{t}$, and 0 otherwise; $\bar{t}$ corresponds to the date of a break; d_t takes 1 at the time of the break, and 0 otherwise; and $d_t^a = 1$ if $t > \bar{t}$, and 0 otherwise. Clearly, d_t^b, d_t and d_t^a are break indicators to qualify the period prior to the break, the time at the break and the period after the break. Under the null hypothesis of unit root, we should have $\omega_1 = \omega_2 = \omega_3 = \omega_4 = 0$. Substituting (2) into (1), we get

$$x_t = \underbrace{\omega_0 + \omega_1 d_t^b + \omega_2 d_t + \omega_3 t + \omega_4 d_t^a}_{c_t} + \beta x_{t-1} + \sum_{k=1}^{p-1} \phi_k \Delta x_{t-k} + e_t, \tag{3}$$

where $\hat{\beta}$ is the coefficient on which the unit root test is based and ω_k for $\omega_k \neq \omega_0$ are deterministic coefficients modifying the level of ω_0 for $\omega_{k-1} \neq 0$. The upper part of Table 3 gives the deterministic coefficients[7] ω_k, while the bottom part of Table 3 provides the number and the time of the breaks for each x_t.

Including single breaks associated with one structural change over the sample period shows that import becomes stationary. In contrast, gasoline price and income are still nonstationary. Since a single break for a period of 11 years captures either the highest or the lowest point, intermediary structural changes are ignored. Given that the minimum time span is 15 months, we count the number and identify the timing of the breaks over the sample period. We identify 3, 4, 5 and 5 breaks in gasoline consumption, imports, gasoline price and income, respectively. The timing and the number of these breaks differ from one variable to another, which suggests that the effects of an event may (or may not) manifest themselves at different times for each of the relevant variables. This also indicates that economic variables differ in the ways they respond to abnormal and structural shocks. These breaks follow a major socioeconomic event. For example, the July 2006-war changed the structure of the four variables, but the change occurred at different months.

On the basis of these break dates, we run a unit root test for each of the inter-periods. We do not report the statistics for these tests, but we could observe that a nonstationary process can be stationary itself in some inter-periods. Specifically, gasoline consumption is stationary in each of the inter-periods, whereas gasoline price, imports and income are nonstationary in one of the inter-periods. Subsequently, we conclude that gasoline price, import and private spending are stationary processes, but with strong shifting drifts.

2.3. Estimating the elasticity of demand

We follow the tradition in the energy economics literature when estimating price and income elasticity (e.g., Eltony, 1994; Baltagi and Griffin, 1997; and Hughes et al., 2008). Our starting point is the partial adjustment model given by

$$q_t = \beta_0 + \beta_l q_{t-l} + \beta_2 p_t + \beta_3 z_t + \varepsilon_t, \tag{4}$$

where β_i measures feedbacks in consumption, l indexes the optimal lag, β_2 measures the price elasticity, β_3 measures the income elasticity, z_t is either y_t or m_t, and ε_t is the error term. The price of gasoline p_t is exogenously determined given that Lebanon is a small-open economy. Lebanon is a price-taker and accordingly p_t cannot cause a simultaneity bias in our estimates. Additional variables related for example to the use and the stock of vehicles could be possibly added to Eq. (4). In the course of our empirical investigation, we did explore this possibility without much success as the data at our disposal are too erratic. Keeping with the tradition in the literature, a positive sign is expected on β_1 as gasoline demand tends to persist from one period to another, a negative sign on β_2 as a price increase reduces quantity demanded, and a positive sign on β_3 indicates that gasoline is a normal good.

However, restricting consumption feedbacks to one period has the potential to bias our estimates. Not only because p_t and z_t are nonstationary processes, but also because the memory of q_t is shorter than what its autocorrelation function suggests. In fact, an attempt to estimate Eq. (4) as such resulted into failure because the error term exhibited serial correlation. With only 11 years of data, setting lags in consumption at 12 months (i.e., using yearly frequency as in most of previous studies) gives biased coefficients since such coefficients are asymptotically inconsistent. Instead,

[7] Reported values of coefficients in Table 3 have no direct economic meaning.

Table 4
Estimates on the demand function of gasoline with single breaks.

	Model 1	Model 2	Model 3	Model 4
$\hat{\beta}_{11}$ of $q_{t-1}d_{t-1}^{qb}$ in (5)	0.040		0.041	
$\hat{\beta}_{12}$ of $q_{t-1}d_{t-1}^{qa}$ in (5)	0.256**		0.269**	
$\hat{\beta}_1$ of q_{t-1} in (4)		0.263**		0.272**
$\hat{\beta}_2$ of q_{t-2} in (4)		0.282**		0.260**
$\hat{\beta}_3$ of q_{t-3} in (4)		0.111		0.070
$\hat{\beta}_2$ of p_t in (4)		−0.258		−0.623**
$\hat{\beta}_{21}$ of $d_t^{pb}p_t$ in (5)	−0.915**		0.028	
$\hat{\beta}_{22}$ of $d_t^{pa}p_t$ in (5)	−0.835**		−0.051	
$\hat{\beta}_3$ of m_t in (4)		0.815**		0.309**
$\hat{\beta}_{31}$ of $d_t^{mb}m_t$ in (5)	0.424**		0.787**	
$\hat{\beta}_{32}$ of $d_t^{ma}m_t$ in (5)	0.414**		0.756**	
$\hat{\beta}_{01}$ of d_t^{qb} in (5)		−3.234**	−3.090**	
$\hat{\beta}_{02}$ of d_t^{qa} in (5)		−2.953**	−3.457**	
Adjusted R-square, $\overline{R}^2$	0.350	0.298	0.360	0.272
Log likelihood function value	−116.313	−122.874	−114.129	−125.742
Ljung–Box statistic, ε_t	16.422	9.465	16.712	13.376
Ljung–Box statistic, ε_t^2	11.793	14.288	8.604	16.771

d_t^{qb} is an indicator for structural change before to a break, d_t^{qa} is an indicator for structural change at and after a break, q_t is gasoline consumption, p_t is gasoline price, m_t is income in terms of import, and ** means significance at 10% or less.

we attempt a quarterly pattern as in Samimi (1995). So, when estimating (1), we lag consumption at 3.

Although, keeping feedback in consumption at lag 1 as in (4) requires another representation of (4). We consider in this paper an approach that uses structural breaks, not only in consumption, but also in price and income. In this way, we obtain a dynamic characterization of short-run elasticities. Moreover, because the dynamics of structural changes are not restricted to breaks in gasoline consumption, short-run elasticities in price and income may be independent of feedbacks in consumption.

Let d_t^b be an indicator before the break for any x_t, and d_t^a be an indicator at and after the break for any x_t, then we estimate the following dynamic model for a single break in x_t,

$$q_t = (\beta_{01}+\beta_{11}q_{t-1})d_t^{qb}+(\beta_{02}+\beta_{12}q_{t-1})d_t^{qa} + \left(\beta_{21}d_t^{pb}+\beta_{22}d_t^{pa}\right)p_t+\left(\beta_{31}d_t^{zb}+\beta_{32}d_t^{za}\right)z_t+\varepsilon_t, \tag{5}$$

where β_{01} and β_{02} identify the consumption level before and after the break, β_{11} and β_{12} are associated with consumption dynamic before and after the break, β_{21} and β_{22} are associated with the price elasticity before and after the break, and β_{31} and β_{32} are associated with the income elasticity before and after the break.

Eq. (5) captures multiple effects of structural breaks on gasoline demand. That is, even though breaks occur at different times, and responses of consumption to variation in price and income may not be simultaneous, multiple breaks that belong to the same socioeconomic event determine altogether the sign and the magnitude of the elasticity coefficient.

Alternatively, we could express the dynamic effect of structural changes in terms of breaks in consumption. However, this is clearly suboptimal. To show why, suppose $d_t^{qb}=0$ then d_t^{pb} and d_t^{zb} will be also 0, which implies that q_t is a function of p_t and z_t when $q_{t-1}\neq 0$. Our functional structure is general because d_t^{bq}, d_t^{bp} and d_t^{bz} are not necessary zero, which implies that q_t is a function of p_t and z_t even though $q_{t-1}=0$. Subsequently, we expect the autocorrelation function of the error term to be uncorrelated. Table 4 reports the estimates of Eqs. (4) and (5).

Table 4 combines the estimates of Eq. (4) given by Models 2 and 4 with the estimates of Eq. (5) given by models 1 and 3. It is worth noting that Eq. (4) is estimated under the view that a quarterly pattern better captures the adjustment of gasoline demand to shocks and news. In contrast, Eq. (5) is estimated under the view that feedbacks in gasoline consumption are short-lived when structural changes are taken into consideration. In fact, the estimates of (4) and (5) are robust to both autocorrelation in residuals (looking at the Ljung–Box statistics of the simple residuals) and heteroskedasticity (looking at the Ljung–Box statistics of the squared residuals). Moreover, feedbacks in gasoline consumption, prices and incomes explain up to 36% of variation in gasoline demand. Since our residuals do not show historical patterns and the adjusted R-square take reasonable values, our estimates cannot be spurious.

While we assumed a quarterly pattern and we could be tempted to use yearly data as it is the case in the literature, the data process proved us wrong, as it takes only two months for gasoline demand to adjust to news and shocks. However, the estimates associated with price elasticities are not significant in Model 2 and 3 even though the coefficients exhibit the correct sign. In contrast, Model 1 and 4 exhibit significant price and income elasticities, which are by their magnitudes higher than short-run elasticities in the literature (see Espey, 1998; and Graham and Glaister, 2002). Focusing on Model 4 for which the short-run price and income elasticity are 0.623 and 0.309, respectively; the long-run price and income elasticity at month 1, 2 and the sum of the three months' are 0.856, 0.842 and 1.564, and 0.425, 0.418 and 0.776, respectively. Comparing these estimates, the long-run elasticities are in line with previous studies. Looking at Model 1, price and income elasticity are 0.915 and 0.424, and 0.835 and 0.414 for structural changes before and after the break, respectively. These price elasticities are higher than in Model 4. However, we refrain from computing long-run price and income elasticity as these estimates are obtained from a subset of the data.

Including breaks in Eq. (4) shortens the adjustment time to news and shocks. However, a single break ignores other occasional breaks giving rise to long life structural changes. Therefore, we extend Eq. (5) with multiple breaks in consumption, price and income. We search for local maxima and minima over three years to identify all the intermediate breaks. Thereafter, we estimate the following model including multiple structural changes to explain gasoline demand,

$$q_t = \sum_{i=1}^{BQ}\beta_{1i}q_{t-1}d_{t-1}^{qi}+\sum_{i=1}^{BP}\beta_{2i}p_t d_t^{pi}+\sum_{i=1}^{BZ}\beta_{3i}z_t d_t^{zi}+\varepsilon_t, \tag{6}$$

where d_t^{xi} is an indicator for structural change; BQ, BP and BZ are the total number of breaks in consumption, price and income, respectively. For example, the three breaks in consumption fall on 2006:07, 2007:01 and 2009:01 giving rise to the following four indicators: d_t^{q1} for $t<2006{:}07$ and 0 otherwise; d_t^{q2} for $t\geq 2006{:}07$ and $t<2007{:}10$, and 0 otherwise; d_t^{q3} for $t\geq 2007{:}10$ and $t<2009{:}01$, and 0 otherwise; and d_t^{q4} for $t\geq 2009{:}01$ and 0 otherwise. The rest of the indicators for the other variables are defined in the same way. With multiple breaks elasticities are clearly time-varying. Table 5 reports the estimates of Eq. (6).

The second part of Table 5, including the Ljung–Box statistics, the adjusted R-square and the log-likelihood function value show that multiple breaks enhance the magnitude of the estimates. The log-likelihood values and the coefficients of determination are higher than those in Table 4. Since the Ljung–Box statistics for simple and squared residuals are uncorrelated up to lag 12, we conclude that the estimates are robust both to autocorrelation and heteroskedasticity. In addition to feedback coefficients and price elasticity coefficients, Table 5 reports also income elasticities in terms of imports (Model 6), and income elasticities in terms of private spending (Model 5).

Table 5
Estimates of the demand function of gasoline with multiple breaks.

	Model 5	Model 6
$\hat{\beta}_{11}$	−0.207	−0.222
$\hat{\beta}_{12}$	0.051	0.119
$\hat{\beta}_{13}$	0.395**	0.480**
$\hat{\beta}_{14}$	−0.077	−0.227
$\hat{\beta}_{21}$	−1.085**	−1.184
$\hat{\beta}_{22}$	−1.280**	−1.219**
$\hat{\beta}_{23}$	−1.339**	−1.124**
$\hat{\beta}_{24}$	−1.435**	−1.181**
$\hat{\beta}_{25}$	−0.735	−0.068
$\hat{\beta}_{26}$	−0.711	−0.194
$\hat{\beta}_{31}$		0.578**
$\hat{\beta}_{32}$		0.668**
$\hat{\beta}_{33}$		0.620**
$\hat{\beta}_{34}$		0.511**
$\hat{\beta}_{35}$		0.631**
$\hat{\beta}_{41}$	1.218**	
$\hat{\beta}_{42}$	1.173**	
$\hat{\beta}_{43}$	1.080**	
$\hat{\beta}_{44}$	0.970**	
$\hat{\beta}_{45}$	0.848**	
$\hat{\beta}_{46}$	0.529**	
Adjusted R-square, $\overline{R}^2$	0.416	0.506
Log likelihood function value	−106.877	−103.976
Ljung–Box statistic, ε_t	16.417	13.613
Ljung–Box statistic, ε_t^2	6.744	8.384

$\hat{\beta}_{1i}$ are related to gasoline consumption, $\hat{\beta}_{2i}$ are related to price elasticity, $\hat{\beta}_{3i}$ are related to income elasticity in terms of imports, $\hat{\beta}_{4i}$ are related to income elasticity in terms of private spending, and ** means significance at 10% or less.

The magnitude of the price and the income elasticities are higher than those in Table 4. This is a direct outcome of including more breaks in Eq. (6). Regardless of the model, Table 5 clearly shows that the demand is price and income elastic. However, the elasticity coefficients are regime-dependent. For example, price elasticities were highest in Model 5 between August 2006 and March 2008, a period encompassing several dramatic events in Lebanon. Contrasting the estimates of Table 5 with those of Table 4, it is clear that the higher the number of breaks, the larger elasticities tend to be. This is understandable as breaks are usually associated with changes in economic policies or turning points in the aftermath of political turmoil.

3. Concluding remarks

Understanding the sensitivity of gasoline demand to changes in price and income is important for economic and environmental reasons. However, in spite of a large number of studies on the subject, the relationship is still puzzling. Gasoline consumption is price and income inelastic in the short-run. A puzzle that should despair environmentalists in their fight against the excessive use of fossil fuels, but should content energy lobbyists and public tax agencies maximizing private interests in the first case and tax revenues in the latter. While the short-run price and income elasticities of gasoline demand have been estimated for many countries, a careful examination of the literature reveals the need to investigate this issue from another perspective. One perspective that is absent in the literature, notwithstanding the relevance of the Lebanese case, is the use of structural breaks in gasoline demand models.

In this respect, our model considers gasoline demand not only as a function of price and income, but also as a function of structural changes in consumption, price and income. Specifically, demand is found to be price inelastic in the short-run in a model including either a single break or without a break. However, both price and income elasticities are higher than what is commonly reported in the literature when multiple breaks are considered. Thus, these elasticities indicate that gasoline demand changes structurally. Therefore, we conjecture that households may rely less on their private vehicles and more on alternative transportation modes (mainly 'service' taxi)[8] as the price of gasoline increases in some inter-periods of structural changes. Furthermore, changes in income seem to impact gasoline consumption more when structural breaks are introduced. This indicates that gasoline is a quasi-luxury good in the Lebanese case.

Finally, our findings have implication on taxation policies. For instance, our elasticity estimates, which are on the higher side, imply that the current flat gasoline tax in effect in Lebanon – while potentially serving an environmental objective in reducing emissions – may not serve well the purposes of policy makers in maximizing government tax revenues. Indeed, policy makers should consider the benefits of adjusting the value of excise tax periodically to reflect changes in lifestyle, income level, and political conditions. However, when excise tax is adjusted upwards its impact will be asymmetric across individuals in different income brackets such that individuals in the lowest income bracket are at disadvantage. This is especially relevant in the Lebanese context, where the use of private cars is widespread, while public transit is underdeveloped. Therefore, the Lebanese government would have to put in place a compensation schedule to offset the negative effect of a tax increase on individuals in the lowest income bracket (see Sterner (2011)). Clearly, there should be a substitutive transportation mode that is accessible and competitive to make such tax increase socially acceptable.

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[8] Affordable shared taxis are known as service in Lebanon, and should not be confused with the regular taxi service.

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