Reconsidering the Time-Series Approach to Estimating Cigarette Price Elasticities when Exploitation of Spatial Variation Appears Unattainable: The Case of Slovakia[1]

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Abstract

Although the current dominant trend in the literature is to analyze cigarette demand through individual consumer microdata, an aggregate country-level approach might be more preferable for countries that legislatively prohibit spatial variation in their prices. This paper, therefore, explores the idea of utilizing the cigarette tax declaration data to estimate a monthly vector error correction model of cigarette prices and volumes taxed in Slovakia for a relatively stable period of Jun 2011 – May 2023. The results are also confronted with alternative estimates using the univariate autoregressive distributed lag model framework. Despite observing volumes of cigarettes stocked by Slovak retailers, after controlling for several exogenous variables, the estimated price elasticity tends to approach magnitudes quite similar to those found in the literature investigating inelastic cigarette demand. Additionally, the paper also provides evidence for measures of tobacco control policy effectiveness regarding the cigarette tax elasticity of volumes stocked by retailers.

Keywords: *cigarette demand, cigarette supply, excise duty*

JEL Classification: D12, D22, H25, I18, L66

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Introduction

The recent wave of cigarette demand analyses based on individual consumer microdata in European countries (see, e.g, Zubović et al., 2019) was not avoided by Slovakia (Jamrich and Pokrivčák, 2018; Lichner and Ostrihoň, 2024; 2025). However, the Slovak legislative environment poses a peculiar challenge from the perspective of microdata analysis. Specifically, the currently in force Act 106/2004 Coll. on Excise Duty on Tobacco Products legally prohibits any spatial variation in cigarette prices across the entire territory, which renders relatively strict assumptions of Deaton's (1997) approach used by Cizmovic et al. (2022), Gligorić et al. (2022), and Vladisavljević et al. (2021) difficult to justify.

On the other hand, the Institute for Financial Policy (IFP) of the Ministry of Finance of the Slovak Republic (MFSR) provides detailed and publicly accessible monthly statistics on tax revenues from tobacco products (see the Appendix, Table A1). A potential hurdle from the perspective of a cigarette demand analysis is that the volumes taxed under Slovak legislation differ from the quantities of cigarettes actually purchased by consumers. However, this complication can also be viewed as an added value of the presented research for policymakers. When, e.g., forecasting tax revenues, the MFSR is likely to be more interested in predicting taxes collected as precisely as possible than the actual cigarette demand.

The paper, therefore, utilizes Slovak cigarette tax declaration data to model cigarette prices and volumes taxed via monthly vector error correction models (VECMs) with exogenous variables and autoregressive distributed lag (ARDL) models. Apart from providing implied own-price and income elasticity estimates, the paper also illustrates the reaction of aforementioned variables by means of impulse response functions (IRFs). The results indicate an own-price elasticity of cigarette volumes taxed not exceeding (or even reaching) unity. Furthermore, the alternative model approaches provide insight into the tax elasticity of cigarette volumes stocked by retailers, which is inelastic in the long run.

The novelty of the approach applied in this paper, compared to other existing time-series cigarette demand analyses, lies in the accommodation of the specificities of the Slovak legal framework. Particularly, isolating cigarette volumes taxed with the combined tax rate, for which the weighted average price can be derived from the data, and accounting for the fact that a considerable portion of the Slovak cigarette market was omitted in the earlier period used in the analysis. The contribution of the paper is, hence, in the following four directions: (i) deriving the monthly weighted average price of cigarettes (WAPC) taxed with combined tax rate for Slovakia using the tax declaration data, which allowed for the application of the multivariate VECM approach; (ii) demonstrating that accounting for the structure of Slovak cigarette market can allow for producing cigarette own-price elasticities similar to those for consumer demand; (iii) providing reference for cigarette tax elasticity, which is scarcely discussed in regard to Slovak tobacco control policy; and (iv) gauging the ability of constructed models to predict volumes of cigarettes taxed.

The rest of the paper is structured as follows. The second section presents the methodology employed. The third section describes the derivation of the monthly WAPC and the additional data used in the analysis. The fourth section presents and discusses the results. Final remarks are provided in the concluding section.

1. Literature Review

When analyzing cigarette demand, it is a common practice nowadays to rely on survey microdata of consumers and apply Deaton's (1997) approach to control for endogeneity in the prices reported by respondents. The approach is based on the assumption that there is a spatial variation in prices, resulting from differences in factors such as transportation costs and the regional demand, which enables the identification of exogenous price changes.

This convenient approach was recently applied in southern European countries, specifically by Cizmovic et al. (2022) for Montenegro, Gligorić et al. (2022) for Bosnia and Herzegovina, and Vladisavljević et al. (2021) for Serbia. For all of these countries, the respective authors found mostly inelastic demand response to cigarette price changes, with the total own-price elasticity of cigarette demand exceeding unity only for low-income households.

However, the microdata analysis becomes slightly less appealing in cases where countries legislatively prohibit spatial variation in cigarette prices, as documented for Slovakia by Lichner and Ostrihoň (2024; 2025). Not only is it difficult to justify Deaton's (1997) approach in such cases, but the authors also report that the minimum tax on cigarettes might have been weak as an instrumental variable. Despite being unable to control for price endogeneity, Lichner and Ostrihoň (2024) obtained estimates of the conditional own-price elasticity of cigarette demand in Slovakia centered at about –0.8. Such estimates were quite consistent with an analogous elasticity at approximately –0.9 provided by Jamrich and Pokryvčák (2018), who used the Heckman sample-selection model. Unlike the previous elastic estimates for Balkan countries, Lichner and Ostrihoň (2025) report elastic demand response to price changes for middle-income households. Since Slovakia was also used in cross-country panel data studies, another own-price elasticity reference can be derived from Kohler et al. (2023), who estimated magnitudes from –0.3 to –0.45 for a panel of European countries and from –0.61 to –1.24 for individual annual samples from that panel.

In parallel with the microdata analyses, a time-series approach to estimating cigarette price elasticities has emerged in the literature, which relies on country-level data. Multiple model frameworks were employed in this regard (Nguyen et al., 2012), of which the most relevant for the presented analysis is the error correction model. This model framework was applied by Hondroyiannis and Papapetrou (1997) to estimate the own-price elasticity of cigarette demand in Greece. The authors initially applied the Johansen and Juselius procedure within a vector autoregressive model. However, they opted for a univariate framework when examining the short-term dynamics akin to the Engel-Granger approach. Furthermore, the authors used the number of cigarettes produced and imported in Greece as a basis for examining cigarette consumption. Similarly, Nikolaou and Velentzas (2001) repeated the analysis of the aforementioned authors for Greece, using a longer time span and real instead of nominal per capita disposable income.

Nguyen et al. (2012) also used the Engel-Granger approach, among several others, to estimate the own-price elasticity of cigarette demand in 11 European countries. The authors utilized "apparent cigarette consumption" for some countries, which they computed as cigarette production plus imports minus exports. Ross et al. (2012) followed suit with a similar analysis for Ukraine, as part of which they confronted regression-based elasticity estimates with point price elasticity estimates derived from annual changes in consumption and prices. Rodríguez-Iglesias et al. (2017) applied the Engel-Granger approach to estimate short- and long-run cigarette demand elasticities in Argentina. Most recently, Ciccarelli et al. (2018) employed, among other approaches, the Engel-Granger methodology to investigate the demand for aggregate tobacco and its various components in Italy. The authors also utilized data derived from cash payments for tobacco products intended for domestic consumption by authorized dealers to monopoly sale warehouses as a proxy for particular tobacco product consumption.

The most relevant development in the field of tobacco control for the presented analysis was the shift to multivariate VECMs, as shown in studies of Martínez et al. (2015) and Marzioni et al. (2023). Martínez et al. (2015) utilized VECM to model cigarette demand, cigarette price, and income in Argentina, while taking into account selected exogenous dummy variables. However, the authors provide the estimates of short-term dynamics only for the cigarette demand. Marzioni et al. (2023), on the other hand, employed the VECM with exogenous variables to explain the interaction between demands for cigarettes and heated tobacco products in Italy. Additionally, the authors also employed the ARDL model as a robustness check of their results.

The ARDL methodology was for the purposes of modeling cigarette demand, already utilized by Mushtaq et al. (2011), who filled the gap in estimating cigarette price elasticity in Pakistan based on "apparent cigarette consumption" (similar to Nguyen et al., 2012). More recently, the same approach was applied to investigate the cigarette demand in Turkey (Yildiz, 2020) as well as to provide additional evidence for Pakistan (Ullah Khan and Shah, 2020, following the framework of Mushtaq et al., 2011).

Employing the extended approach of the non-linear autoregressive distributed lag model, Martín Alvarez et al. (2020) estimated price and income elasticities of cigarette demand in Spain. Most recently, Seleka and Agang (2025) estimated price and income elasticities of tobacco demand in Botswana using the ARDL approach.

2. Methodology

Similar to Hondroyiannis and Papapetrou (1997), Mushtaq et al. (2011), Nguyen et al. (2012), and Ciccarelli et al. (2018), the analysis in this paper relies on cigarette data based on the supply side rather than the demand side. Although there is a long tradition in utilizing such data for demand analyses, an additional theoretical underpinning supporting this approach is the work of Muhammad and Hossen (2025), which treats imported products from the perspective of intertemporal utility maximization. Since almost all cigarettes sold in Slovakia are not domestically produced (Hudcovský and Morvay, 2024b), this approach was considered appropriate.

Because of this, a standard model of cigarette demand based on a log-log specification was employed. After investigating the stationarity (see the Appendix, Table A2) of the underlying time series, the VECM framework, akin to Martínez et al. (2015) was applied. Similar to the aforementioned authors and Marzioni et al. (2023), additional independent variables were included in the estimated VECMs as unmodeled exogenous variables. To be more precise in this regard, the presentation of Johansen's methodology provided by Hjalmarsson and Österholm (2010), expanded following EViews (2025), was adopted to detail the approach:

$$\Delta \mathrm{y}_t = \mu + \mathbf{\Pi}\mathrm{y}_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta \mathrm{y}_{t-1} + \mathbf{B}\mathrm{z}_t + \varepsilon_t \tag{1}$$

In equation (1), y_t is the $n \times 1$ vector of endogenous variables, z_t is the $m \times 1$ vector of exogenous (deterministic) variables, and ε_t is the $n \times 1$ vector of innovations (shocks). $\mathbf{\Pi}$ is the $n \times n$ coefficient matrix which can be factorized into $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $n \times r$ matrices of adjusting speed parameters and cointegrating vectors, respectively. Additionally, $\mathbf{\Gamma}_i$ are $n \times n$ coefficient matrices at different lags of differenced endogenous variables, and $\mathbf{B}$ is the $n \times m$ coefficient matrix of exogenous variables. In this case, the endogenous variables y_t are cointegrated among themselves but not with the exogenous variables z_t, i.e., the exogenous variables z_t affect only the short-run dynamics. However, the methodology can be extended to allow for exogenous variables z_t to be cointegrated with endogenous variables y_t, i.e., for the exogenous variables z_t to affect the cointegrating vectors themselves, as follows:

$$\Delta \mathrm{y}_t = \mu + \boldsymbol{\alpha}\left(\boldsymbol{\beta}'\mathrm{y}_{t-1} + \boldsymbol{\eta}'\mathrm{z}_t\right) + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \mathrm{y}_{t-1} + \varepsilon_t \tag{2}$$

In equation (2), $\boldsymbol{\eta}$ is the $m \times r$ coefficient matrix, where r is the number of cointegrating vectors.

Hence, the vector of endogenous variables consisted of the volume of cigarettes taxed per capita (q) and the real weighted average price of one cigarette in 2015 prices (p): $\mathrm{y}_t = \left(q_t \quad p_t\right)'$. In the case of both variables, cigarettes taxed with the minimum tax rate were disregarded.

Allowing for two endogenous variables inherently complicates the interpretation of the cointegrating vector, as it may depict the long-run equilibrium relationship between the demanded quantities and the quantities supplied. Therefore, it is henceforth interpreted as a reduced form of both supply and demand equations, similar to the framework of Cotti et al. (2022).

To better distinguish the endogenous variation, the changes to real specific part of the tax on cigarettes in 2015 prices (henceforth referred to as "changes to constant specific duty", assumed to explain the short-run dynamics) and an alternative definition of changes to nominal specific duty deflated by 2015 prices in the initial year (henceforth referred to as "changes to current specific duty", assumed to explain the short-run dynamics) were used as exogenous variables z_t. Additionally, the real monthly wages in 2015 prices and the real heated tobacco product (HTP) duties[2] in 2015 prices (both considered to affect only the cointegration vector) were used as additional exogenous variables.

Since the specific duty changes are enforced by law, they are announced sufficiently in advance[3] before they become effective. Because of this, a one-month lead of changes to current and constant specific duty were also used as exogenous variables affecting only the short-run dynamics, serving as a means to explain

[2] Since the tax declaration data for HTP duty were available only from May 2017, a quantity of 10^(–4) was imputed for missing values prior to this date to allow for logarithmic transformation.

[3] E.g., Hudcovský and Morvay (2024a) report that upcoming changes to the cigarette excise duty until 2028 were already known in 2024.

potential stockpiling by retailers before the change took effect (similar development in cigarette sales for Ukraine was observed by Ross et al., 2012).

The cointegration of p and q was assessed using the Johansen test and the assumption that the cointegrating relationship does include a constant while the short-run dynamics do not. However, due to q potentially being stationary (see the Appendix, Figure A1), additional tests for near-integrated time series suggested by Hjalmarsson and Österholm (2010) were employed. The stability of the cointegrating vector was evaluated by investigating an analogous specification estimated as a standard univariate linear model via least squares using multiple variants of the Bai and Perron (1998) test. All of the performed tests are at a significance level of $\alpha = 0.05$.

The ARDL model framework[4] was used as an alternative method of statistical verification of the existence of a cointegrating relationship. To this end, Pesaran et al. (2001) bounds testing approach was employed, which was developed for assessing cointegrating relationships in cases when it is not certain that all regressors are stationary or integrated of order one. Similar to the application of the Johansen test for VECMs, a restricted constant was assumed in the cointegrating vector for all ARDL models. Heteroskedastic-autocorrelation consistent standard errors based on Bartlett kernel and Newey-West bandwidth selection relying on Akaike information criterion were used in all estimated ARDL models. The stability of each estimated ARDL model was tested by the cumulative sum control chart (CUSUM) and the recursive residual plot. The number of breaches of the respective 5% control bands is indicated for each model.

Additionally, both in-sample and out-of-sample forecasts[5] were performed for 6 periods before (Dec 2022 – May 2023) and after the end (Jun 2023 – Nov 2023) of the analyzed sample, respectively. The predictive ability of each model in this regard was evaluated by mean absolute percentage error (MAPE), with suffix "in" for in-sample and suffix "out" for out-of-sample performance.

3. Data

Utilizing the IFP MFSR data on tax returns from tobacco products, the WAPC for sticks taxed with the combined tax rate was calculated by expanding on the exemplary approach for cigarette tax computation provided by the Financial Administration of the Slovak Republic (FASR).[6]

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[4] The extension of the VECM analysis by the ARDL model framework was suggested by an anonymous reviewer.

[5] The assessment of the predictive abilities of each estimated model was suggested by an anonymous reviewer.

$$T_t^c = SD_t + AVT_t P_t \tag{3}$$

In equation (3), T_t^c represents the total combined tax rate levied on a single cigarette stick, SD_t is the specific part of the tax, AVT_t is the percentage part of the tax rate (or more colloquially, "ad valorem" tax rate), and P_t is the price of a single cigarette stick, all observed in month t. Sticks taxed with a combined tax rate are those for which the total tax collected exceeds the minimum tax rate on cigarettes. Moving from an arbitrary stick to averages by substituting T_t^c with the average combined tax rate levied on cigarettes taxed only with the combined tax rate ($\overline{T}_t^c$) in equation (3), one can rearrange the terms to derive the level of WAPC for cigarettes taxed only with the combined tax rate ($\overline{P}_t^c$), which is represented as the following equation (4):

$$\overline{P}_t^c = \left(\overline{T}_t^c - SD_t\right) / AVT_t \tag{4}$$

Although in equation (4) $\overline{P}_t^c$ is a function of SD_t, this step is not expected to introduce endogeneity between SD_t and $\overline{P}_t^c$, for the purposes when specific duty is used as an exogenous variable in VECM and ARDL settings. As was mentioned before, the excise duty on cigarettes is determined in advance before it becomes effective, i.e., the tax rate does not respond to current developments in the WAPC. Furthermore, even if the $\overline{T}_t^c$ is misreported, the nominal tax rate is fixed across an extended period. Therefore, there should not be any errors in SD_t, which can be correlated with errors in $\overline{P}_t^c$. Lastly, taxes are well established as instruments for purposes of cigarette demand analyses (see, e.g., Cheng and Estrada, 2020; Cotti et al., 2022).

Analogously, if one substitutes T_t^c in equation (3) with the minimum tax rate on cigarettes ($T_t^{min} \lim\limits_{x \to \infty}$), the minimum price at which sticks are taxed with the combined tax rate (P_t^{min}) can be obtained, as presented in equation (5) below:

$$P_t^{min} = \left(T_t^{min} - SD_t\right) / AVT_t \tag{5}$$

Consequently, the price of sticks taxed with the combined tax rate is higher than the price of sticks taxed with the minimum tax rate, which implies that

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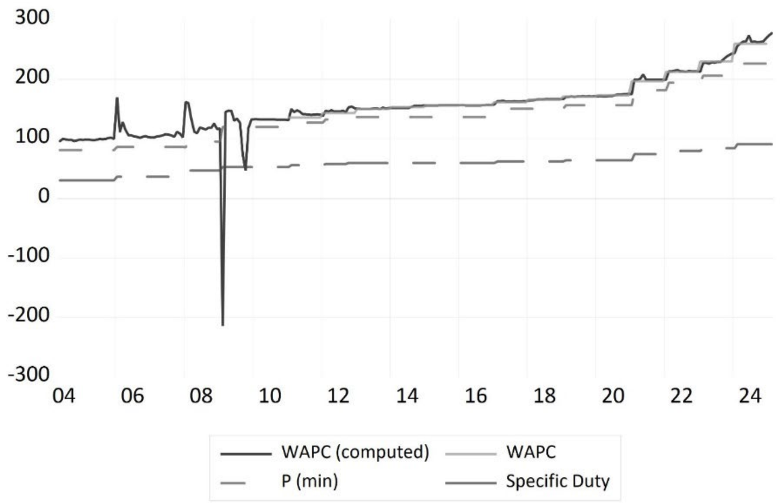
[6] For more details, see <https://www.financnasprava.sk/sk/obcania/dane/spotrebne-dane/spotrebne-dane-obacaia-tabak>. Alternatively, the equation for public revenue from taxation, as presented by Ciccarelli et al. (2018), can be used as a theoretical underpinning for the calculation of the WAPC.

the WAPC for sticks taxed with the combined tax rate is higher than the WAPC for all sticks. The differences between these WAPCs are discernible in Figure 1, below, along with the levels of the minimum price and the specific part of the tax (per 1,000 cigarettes).

F i g u r e 1

Computed WAPC for Sticks Taxed with the Combined Tax Rate, WAPC for All Sticks, the Minimum Price, and the Specific Part of the Tax



Note: "WAPC" / "WAPC (computed)" stands for annual weighted average price of all cigarettes / computed monthly weighted average price cigarettes taxed with the combined tax rate, "P (min)" is the minimum price at which sticks are taxed with the combined tax rate, and "Specific Duty" stands for the specific part of the tax on cigarettes. All quantities are enumerated in current EUR per 1,000 cigarettes. The data are presented over the period May 2004 – Feb 2025.

Source: Author's own calculations based on equations (4) and (5), the IFP MFSR, and the FASR (for further reference to the data sources see Table A1 in the Appendix and footnote 10 on page 420).

As shown in Figure 1, the WAPC for sticks taxed with the combined tax rate calculated in this manner is quite volatile, with notable dips below the minimum price level, particularly in February, September, and October of 2009.[7] However, the quantities of cigarettes taxed are even more volatile before this period as can be seen in Figure 2, below. The figure also shows that the share of cigarettes taxed with the minimum tax rate is quite substantial and becomes negligible towards the latter half of 2013.[8]

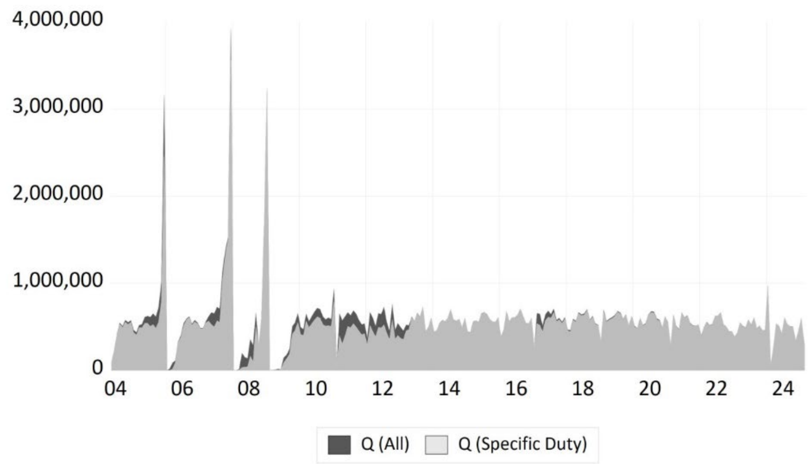
[7] Based on a consultation with employees of the IFP MFSR, there may have been errors in tax declarations reported in February 2009, which resulted in the negative value of the computed WAPC.

[8] Between February 2017 and August 2017, the share of cigarettes taxed with the minimum tax rate on the total volume of cigarettes taxed also exceeded 5%.

In terms of reliability, the tax declaration data ought to be more precise than, e.g., self-reported survey data, as misreporting them could be considered tax evasion. Nevertheless, a temporal discrepancy between the volumes of cigarettes that are monthly stocked as inventories and sold to consumers is possible. By comparison with the cigarette consumption data for 2014 – 2018 on an annual level, the tax declaration data are within ±2% of the estimated cigarettes consumed[9] in Slovakia.

F i g u r e 2

Comparison of the No. of Sticks Taxed with the Combined Tax Rate with the Total No. of Sticks Taxed



Note: "Q (All)" stands for the total number of cigarettes taxed in observed month (sum of cigarettes taxed only with combined tax rate and cigarettes taxed with minimum tax rate) and "Q (Specific Duty)" stands for the number of cigarettes for which specific duty is applicable in the observed month (cigarettes taxed only with combined tax rate). All quantities are enumerated in 1,000 sticks. The data are presented over the period May 2004 – Feb 2025.

Source: The IFP MFSR (for more details see Table A1 in the Appendix).

Additional data on the Harmonized Index of Consumer Prices (HICP) for all items, as well as on average monthly wages, were sourced from the National Bank of Slovakia (NBS). Information regarding monthly population levels was available from the Statistical Office of the Slovak Republic (SOSR). The annual WAPC was acquired from the FASR.[10]

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[9] KPMG EU flows model 2014 – 2018, see <https://public.tableau.com/views/CountryReport-TOUPLOAD/CountryOverview?%3Aembed=y&%3AshowVizHome=no&%3Adis-play_count=yes&%3Atoolbar=no#3>.

[10] For more details, see <https://www.financnasprava.sk/sk/danovi-a-colni-specialisti/dane/spo-trebne-dane/dolezite-informacie/_1/dTypAVelkost/g>.

Moreover, the changes to the value-added tax and ad valorem tax on cigarettes were drawn from the relevant legislation.[11] Further description of logarithmic transformations and seasonal adjustments of the used variables is provided in the Appendix, Table A1.

4. Results

4.1. The VECM Framework

The results of the estimated VECMs are presented in Table 1 below. Due to the unavailability of historical monthly wages, the sample for the main results was restricted to the more recent period. An advantage was taken of the fact that both the value-added tax and the cigarette ad valorem tax remained unchanged from Feb 2011 to Jan 2024. To avoid risking any potential changes resulting from lingering retailer adjustments to these policy shifts, prevalent majority of the presented models was estimated on a sample Jun 2011 – May 2023. The initial model (1) in Table 1 is the only exception from this rule, as a representative of the case when all available observations from May 2004 to February 2025 were used, instead. In this case, the minimum price was imputed for the WAPC for observations of WAPC below the minimum price (see Figure 1). Based on the Akaike information criterion, three lags were considered necessary to account for short-run dynamics in cases of specifications without other exogenous variables.

The differences between models (1) and (2) help to illustrate that although the analysis based on the restricted sample (Jun 2011 – May 2023) is not able to confirm existence of cointegrating relationship between p and q as shown by corresponding Johansen test or Hjalmarsson and Österholm test statistics, the extended sample (May 2004 – Feb 2025) corroborates the existence of such relationship for all of the mentioned tests.

Regardless of the sample or model used, the deviations of q from the cointegrating vector are relatively quickly corrected. With an estimated adjusting speed of approximately –0.4 for model (1), the half-life of the deviation is approximately 41 days.[12]

Additionally, the speed is highly statistically significant in all estimated specifications. The same does not hold for p, regarding which the adjusting speed for model (1) is 0.013, i.e., p is diverging from the cointegrating vector.

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[11] For more details, see <https://www.slov-lex.sk/ezbierky/pravne-predpisy/SK/ZZ/2004/222/ and https://www.slov-lex.sk/ezbierky/pravne-predpisy/SK/ZZ/2004/106/>.

[12] Employing the formula used by Boitani and Dragomirescu-Gaina (2023), the approximate half-life of deviation in q from the cointegrating vector is $\ln(0.5)/\ln(1\text{-}0.4) = 1.357$ months.

T a b l e 1

Cigarette Price Elasticity Estimates for Slovakia Based on VECMs

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Cointegration vector	*p*	4.321	0.595	–0.534	0.37	–0.547	1.391	–0.768	2.075	–0.686
	p (t)	[2.586]	[1.009]	[–2.591]	[0.622]	[–2.628]	[1.451]	[–2.454]	[1.875]	[–2.125]
	Wage						–0.103	0.259	–1.026	0.079
	Wage (t)						[–0.19]	[1.421]	[–1.136]	[0.278]
	HTP Duty								0.01	0.002
	HTP Duty (t)								[1.138]	[0.789]
Short-run dynamics in:	*q* – Adjusting speed	–0.409	–0.319	–1.025	–0.306	–1.002	–0.345	–1.228	–0.285	–1.208
	q – Adjusting speed (t)	[–5.359]	[–2.765]	[–5.682]	[–2.721]	[–5.72]	[–3.47]	[–7.226]	[–3.347]	[–7.275]
	q – Spec. Duty				–2.411	–2.361	–1.163	–1.397	–1.096	–1.366
	q – Spec. Duty (t)				[–2.798]	[–2.98]	[–1.794]	[–2.439]	[–1.692]	[–2.393]
	q – Spec. Duty(+1)						–0.757	–0.882	–0.676	–0.836
	q – Spec. Duty(+1) (t)						[–1.173]	[–1.55]	[–1.049]	[–1.473]
	p – Adjusting speed	0.013	0.009	–0.01	0.001	–0.018	0.018	–0.009	0.018	–0.006
	p – Adjusting speed (t)	[3.904]	[0.81]	[–0.507]	[0.09]	[–1.425]	[2.321]	[–0.565]	[2.655]	[–0.386]
	p – Spec. Duty				0.802	0.805	0.481	0.457	0.48	0.459
	p – Spec. Duty (t)				[14.284]	[14.467]	[9.42]	[8.819]	[9.496]	[8.852]
	p – Spec. Duty(+1)						0.003	–0.017	0.001	–0.016
	p – Spec. Duty(+1) (t)						[0.054]	[–0.329]	[0.013]	[–0.305]
	Str. Br. Dum.	No	No	Yes	No	Yes	No	Yes	No	Yes
	No. of Inc. Obs.	245	140	140	140	140	141	141	141	141
	No. of Coef.	17	17	18	19	20	18	19	19	20
	q – R-sq.	0.356	0.597	0.657	0.62	0.679	0.583	0.673	0.58	0.674
	p – R--sq.	0.247	0.085	0.082	0.64	0.645	0.421	0.399	0.428	0.399
	q – Adjusted R-sq.	0.339	0.579	0.642	0.6	0.662	0.564	0.658	0.562	0.659
	p – Adjusted R-sq.	0.228	0.044	0.041	0.62	0.626	0.395	0.373	0.403	0.372
	MAPE (in)	8.685	0.976	1.571	1.188	1.911	1.017	1.304	1.041	1.396
	MAPE (out)		2.411	3.117	2.142	2.879	2.795	3.664	2.69	3.659
	Lag Exclusion Test (p)	0	0.008	0.151	0.002	0.036	0	0.008	0	0.008
	Ljung-Box Q-Test (p)	0.253	0.697	0.856	0.09	0.224	0.140	0.550	0.142	0.580
	Rao F-test of AC(1) (p)	0.425	0.252	0.685	0.004	0.021	0.005	0.781	0.004	0.749
	Johansen Max. Eig.	29.482	8.427							
	Johansen Trace	32.703	12.937							
	Rest. B(1,1) = 0 (p)	0.013	0.492							
	Rest. B(1,2) = 0 (p)	0	0.048							

Note: The first part of the table presents selected parameters of the estimated cointegrating vectors normalized to *q*, which was, for brevity, omitted from the table. These parameters were multiplied by –1 to mimic the sign that the parameters would have obtained on the right-hand side of the long-run equilibrium relationship. Corresponding t-statistics are presented in square brackets below the estimated parameters. The second part of the table shows the estimated parameters for the short-run dynamics associated with the corresponding endogenous variable, denoted by "*p*"/"*q*" at the beginning of the respective row. The lead in (announced) specific duty is denoted as "Spec. Duty(+1)". Alternative definition (see Methodology) of this and the contemporary specific duty variable were used in models (6) – (9). The last part of the table presents the results of statistical tests, along with additional statistics. "Str. Br. Dum." marks whether the structural break dummy, obtaining "1" for period 2011M06 – 2013M04, and "0" otherwise, was included among exogenous variables affecting the cointegrating vector. The final two rows present the p-values of the corresponding likelihood ratio tests for binding restrictions of B(1,1) = 0; B(1,2) = 1 and B(1,1) = 1; B(1,2) = 0, respectively. Models (6) – (9) are based on two lags of endogenous variables instead of three, which was the default used for other models.

Source: Author's own estimates based on the data specified in Table A1 in the Appendix.

However, when the structural break is accounted for using the dummy variable, the adjustment of *p* is observable but statistically insignificant. The insignificance of the adjustment parameter may indicate the potential exogeneity of *p*, since the price does not appear to adjust to the estimated long-run equilibrium. On the other hand, the *q* is over-adjusting the deviations when the structural break dummy is included in the specification.

Regarding the estimated elasticities, if the structural break is not taken into account by the means of the structural break dummy (models 1, 2, 4, 6, and 8), the own-price elasticity of cigarette volumes taxed, implied by the cointegrating vector, is positive, ranging from 0.37 to 4.321. Analogous income elasticity is statistically insignificant and negative, in both cases contradicting what one would expect based on relevant literature. Although the positive parameter assigned to *p* can be potentially explained by the long-run relationship capturing the factors driving cigarette supply, it is also possible that the estimated parameters are biased by the structural break.

The dummy variable was identified by multiple structural break tests (for more details see the Methodology section), and the period of the break (May 2013) coincides with the shift in the cigarette market to feature only a negligible share (less than 5%) of cigarettes taxed with the minimum tax rate (see Figure 2). This surge of quantities, which were not previously accounted for by models without the dummy variable, may bias the estimates.

Models (3), (5), (7), and (9), which include the structural break dummy, are, therefore, considered more reliable. For estimates of all of these models, the implied own-price elasticity is negative and statistically significant, ranging from –0.534 to –0.768, which is more in line with the literature on cigarette demand (see, e.g., Kohler et al., 2023). Similarly, the income elasticity turns from negative to positive after the structural break is accounted for, although the estimates are still statistically insignificant.

Regardless of the inclusion of the structural break dummy, the HTP cross-price elasticity of volumes of cigarettes taxed remains statistically insignificant and negligibly positive, potentially hinting at a relationship between cigarettes and HTPs as substitutes.

Regarding the short-run dynamics, there is strong evidence that a specific duty on cigarettes affects both *p* and *q* (based on the high t-statistic as well as more than doubling the explained variation of *p*), irrespective of whether the structural breaks are being accounted for.

The effect of contemporaneous percentage change in specific duty on percentage changes in *q*, to some degree, gauges the short-run tax elasticity of cigarette volumes stocked by retailers. This estimate is, however, imprecise, given that the contemporaneous percentage change in *p* is not featured in the short-run dynamics equation for *q*. The estimated effect should therefore be interpreted as a reduced form parameter capturing both the contemporaneous own-price elasticity. The estimated effect at values ranging from –2.4 to –1.1, depending on whether constant or current change was considered, is, nevertheless, rather high. A potential explanation might be a precautionary reaction by retailers to changes in cigarette prices, aimed at avoiding potential losses from excessive response in the cigarette end-user demand.

Regarding the retailers' behavior, there is a lack of evidence that announced changes in the current specific duty are the cause of the stockpiling of q. The estimated parameters are negative and statistically insignificant.

In terms of statistical verification, models (4) – (6) and (8) are considered unreliable due to failing the Rao F-test for serial correlation of order 1. This was an issue, which could not have been resolved even with the inclusion of an additional lag of the endogenous variables.

4.2. Alternative Estimates Based on The ARDL Framework

To provide additional evidence regarding the existence of a cointegrating relationship between p and q, alternative ARDL models are presented in Table 2, below. Similar to the results presented in Table 1, the effect of the structural break was gauged by the comparison of models with (models 1, 3, and 5) and without (models 2, 4, and 6) the dummy variable.

Mimicking the VECM results before, the Pesaran et al. (2001) Bounds test confirms cointegration for all of the models with the structural break dummy, while rejecting it for the majority of models without the dummy. Nevertheless, among models without the dummy, model (5) confirms cointegration, potentially indicating that once additional variables affecting cigarette demand are accounted for, the cointegration relationship becomes discernible even if the structural break is left untreated.

However, among the models presented in Table 2, only models (1) and (2) appear reliable, as these are the only models that satisfy both applied tests for serial correlation. Regarding the overall stability, all presented models are within the CUSUM bands, but there are several breaches of the Recursive Residuals bands for each model.

In terms of estimated elasticities, the obtained estimates of implied own-price elasticities of cigarette volumes taxed are all negative, ranging from –0.159 to –0.963. As was mentioned before, among these, the estimate of model (2) at –0.611 appears to be most reliable, given the results of the Bounds test and serial correlation tests. These results, therefore, corroborate the notion that cigarette demand is being observed through volumes taxed (although indirectly, with retailers in the position of cigarette end-user intermediaries).

On the other hand, the income elasticity based on the parameter associated with the wage variable is less stable and flips from positive to negative, once the HTP duty is taken into account. The cross-price elasticity based on HTP duty is again negligible and positive. Nevertheless, all of the parameters associated with the wage and HTP duty are statistically insignificant.

Table 2

Cigarette Price Elasticity Estimates for Slovakia Based on ARDL Models

	(1)	(2)	(3)	(4)	(5)	(6)
p	–0.159	–0.611	–0.945	–0.963	–0.659	–0.819
p (t)	[–0.239]	[–3.449]	[–1.061]	[–2.092]	[–0.514]	[–1.448]
Wage			0.405	0.239	–0.222	–0.086
Wage (t)			[0.771]	[0.887]	[–0.163]	[–0.149]
HTP Duty					0.007	0.003
HTP Duty (t)					[0.748]	[0.823]
q – Adjusting speed	–0.249	–0.887	–0.367	–0.833	–0.342	–0.807
q – Adjusting speed (t)	[–1.783]	[–4.61]	[–2.525]	[–4.775]	[–2.629]	[–5.11]
Str. Br. Dum.	No	Yes	No	Yes	No	Yes
No. of Inc. Obs.	145	146	138	138	138	138
No. of Coef.	8	8	20	21	21	22
q – R-sq.	0.306	0.378	0.461	0.515	0.463	0.518
q – Adjusted R-sq.	0.27	0.347	0.374	0.432	0.371	0.431
MAPE (in)	1.045	1.566	1.125	1.34	1.121	1.339
MAPE (out)	1.984	2.407	1.608	1.875	1.646	1.926
Bounds test	1.826	7.73	2.98	6.605	3.926	6.745
I(0) Crit. Val.	3.62	3.62	3.1	3.1	2.79	2.79
I(1) Crit. Val.	4.16	4.16	3.87	3.87	3.67	3.67
Ljung-Box Q-Test (p)	0.809	0.876	0.204	0.356	0.175	0.305
Breusch-Godfrey of AC(1) (p)	0.144	0.486	0.006	0.037	0.004	0.023
CUSUM	0	0	0	0	0	0
Recursive Residuals	8	8	7	11	6	10

Note: The first part of the table presents selected parameters of the estimated cointegrating vectors. Corresponding t-statistics are presented in square brackets below the estimated parameters. The second part of the table shows the adjustment parameter for the short-run dynamics associated with the endogenous variable *q*. "Str. Br. Dum." marks whether the structural break dummy obtaining "1" for period 2011M06 – 2013M04, and "0" otherwise, was included among exogenous variables affecting the cointegrating vector.

Source: Author's own estimates based on the data specified in Table A1 in the Appendix.

Although the VECM framework presented in Table 1 has not confirmed that *p* would respond to the long-run equilibrium relationship, as a robustness check to unlikely endogeneity of *p*, the ARDL models were re-estimated using specific duty on cigarettes in constant prices as a proxy for *p*. In this manner, the models presented in Table 3 can be considered either a reduced form of the case when the specific duty would be used as an instrument for *p* or they can be interpreted as models estimating the long-run tax elasticity of *q*.

Regardless, the models presented in Table 3 appear to be more stable than models utilizing *p*, judging by the same or lower number of breaches of Recursive Residual bands, with the exception of model (6), which has a higher number of breaches than its analogue in Table 2. Furthermore, all models except for model (1) confirm cointegration among included variables, and all models except for model (5) pass the used serial correlation tests.

Table 3

Tax Elasticity Estimates for Slovakia Based on ARDL Models

	(1)	(2)	(3)	(4)	(5)	(6)
Spec. Duty	–0.365	–0.725	–0.87	–0.74	–0.704	–0.548
Spec. Duty (t)	[–0.494]	[–4.468]	[–1.347]	[–1.875]	[–0.886]	[–1.077]
Wage			0.179	–0.015	–0.24	–0.376
Wage (t)			[0.56]	[–0.085]	[–0.292]	[–0.844]
HTP Duty					0.005	0.004
HTP Duty (t)					[0.67]	[0.981]
q – Adjusting speed	–0.246	–0.938	–0.368	–0.816	–0.349	–0.716
q – Adjusting speed (t)	[–1.923]	[–3.818]	[–3.409]	[–3.841]	[–3.559]	[–4.314]
Str. Br. Dum.	No	Yes	No	Yes	No	Yes
No. of Inc. Obs.	145	147	138	138	138	138
No. of Coef.	8	7	20	20	21	22
q – R-sq.	0.349	0.404	0.496	0.531	0.497	0.539
q – Adjusted R-sq.	0.316	0.378	0.415	0.455	0.411	0.456
MAPE (in)	1.086	1.986	1.104	1.284	1.102	1.303
MAPE (out)	1.916	2.798	2.075	2.409	2.073	2.381
Bounds test	1.634	7.589	4.222	5.908	4.271	4.962
I(0) Crit. Val.	3.62	3.62	3.1	3.1	2.79	2.79
I(1) Crit. Val.	4.16	4.16	3.87	3.87	3.67	3.67
Ljung-Box Q-Test (p)	0.757	0.748	0.332	0.39	0.296	0.474
Breusch-Godfrey of AC(1) (p)	0.295	0.399	0.06	0.099	0.043	0.145
CUSUM	0	0	0	0	0	0
Recursive Residuals	8	8	5	10	6	13

Note: The first part of the table presents selected parameters of the estimated cointegrating vectors. Corresponding t-statistics are presented in square brackets below the estimated parameters. The second part of the table shows the adjustment parameter for the short-run dynamics associated with the endogenous variable q. "Str. Br. Dum." marks whether the structural break dummy obtaining "1" for period 2011M06 – 2013M04, and "0" otherwise, was included among exogenous variables affecting the cointegrating vector.

Source: Author's own estimates based on the data specified in Table A1 in the Appendix.

In terms of elasticities, the magnitudes of tax elasticity of q presented in Table 3 range from –0.365 to –0.870 and are, thus, within the range for price elasticities presented in Table 2. However, only the tax elasticity based on model (2) in Table 3 is statistically significant. Similar to the results in Table 2, all of the wage and HTP duty elasticities are statistically insignificant.

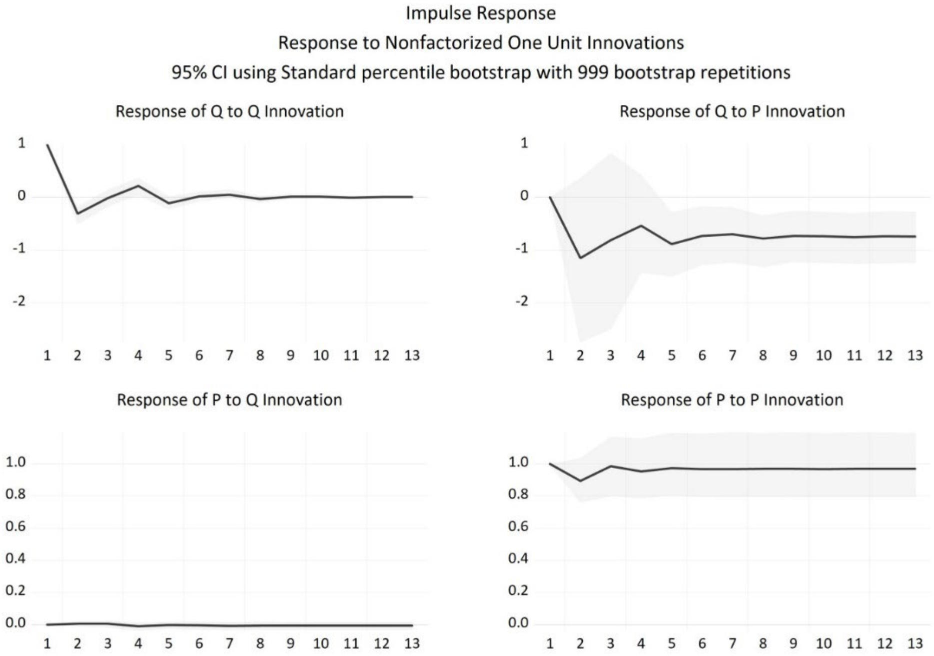
4.3. Impulse Response Function

Based on the validation provided by the ARDL framework, the VECMs accounting for the structural break were considered more appropriately specified as they consistently produce negative own-price elasticities of q. Among these, model (7) produces the best in-sample predictions, with only 1.3% mean absolute percentage error. This performance surpasses even analogous ARDL models that also include the structural break dummy.

Therefore, this model was used to generate impulse response functions (IRFs), which are presented in Figure 3, below.

F i g u r e 3

Impulse Response Functions for One Unit Innovation to Endogenous Variables Based on Model (7) in Table 1



Source: Author's own estimates based on the data specified in Table A1 in the Appendix.

Based on the results presented in Figure 3, an innovation in *p* appears to be rather persistent, showing only small oscillations in immediately succeeding periods. Far more interesting is the response of *q* to the innovation in *p*, which immediately shows a rather dramatic reaction in the following period. However, this response is statistically insignificant, which may potentially indicate different responses to similar situations in the past due to varying levels of cigarette stocks by the retailers. After subsequent adjustment, a statistically significant deviation from pre-shock levels, with a mean of approximately –0.74, emerges in the fifth period and remains significant one year after the shock.

The IRFs, in this regard, corroborate the notion that the long-run own-price elasticity of cigarette volumes taxed might not be that distant from unity. Such elasticity magnitude would be in line with previous estimates of cigarette demand elasticity for Slovakia, obtained using microdata (Jamrich and Pokrivčák, 2018; Lichner and Ostrihoň, 2024; 2025). The relatively wide confidence band of IRFs would cover even smaller estimates of own-price elasticity for cross-country analysis featuring Slovakia by Kohler et al. (2023).

Regarding policy implications, the lack of statistically significant evidence of real wages affecting the volumes of cigarettes taxed may serve as an argument against repeatedly raising the tax rate on cigarettes. However, one should bear in mind that the observed results are for volumes of cigarettes taxed and not cigarette demand itself. Furthermore, a model assigning a positive effect to real wages outperforms models disregarding real wages in terms of the in-sample mean absolute percentage error statistic.

Conclusions

The results of the analysis of monthly cigarette tax revenue data appear to be favorable towards the notion that Slovak own-price cigarette demand elasticities can be estimated using volumes of cigarettes stocked by retailers. Presented approaches might be applicable and attractive for researchers and policymakers in other countries that, similarly to Slovakia, prohibit spatial variation in cigarette prices and have access to quantities at the retail level. Nevertheless, the results also suggest that there are key exogenous factors influencing the volumes and prices of cigarettes taxed in Slovakia. Namely, wages, excise duties on cigarettes, alternatives to smoking such as HTPs, as well as the proportion of cigarettes taxed with the minimum tax rate.

Constructed vector error correction models and autoregressive distributed lag models are consistent with the expectations based on the cigarette demand literature. The results are in the vicinity of previously obtained own-price elasticity estimates for Slovakia based on microdata and cross-country analysis. However, it is important to note that the cointegration relationship essentially falls apart when at least the fact that only a share of the cigarette market is being taxed with the combined tax rate is not taken into account. Besides the estimates of cigarette price elasticity having policy implications in themselves, the model frameworks employed may be more useful than any single parameter obtained, as the in-sample and out-of-sample forecasts exhibit mean absolute percentage errors at approximately 1 – 2% and 1.5 – 3%, respectively. With that said, the results for the short- (which is elastic) and the long-run tax elasticity of volumes stocked by retailers (which is inelastic) may prove useful when setting future tax rates, as such parameter estimates are rarely available for Slovakia.

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Appendix

Table A1

Description of Used Variables

Variable	Description
p	Real weighted average price of cigarettes (WAPC) in 2015 prices. The nominal WAPC was derived following equation (4) using total tax levied, volumes of cigarettes taxed with the combined tax rate, specific part of the tax on cigarettes, and ad valorem tax rate on cigarettes, all publicly available from the IFP MFSR Tax returns – tobacco products database (source: <https://www.mfsr.sk/en/finance/institute-financial-policy/economy-statistics/taxes-contributions/>) and from the corresponding legislation (see footnote 11 on page 421). Subsequently, the nominal WAPC were adjusted using monthly HICP deflators for all products, which are publicly available from the NBS Macroeconomic database (source: <https://nbs.sk/statisticke-udaje/vybrane-makroekonomicke-ukazovatele/makroekonomicka-databaza/?timeSeriesId=%5B%221925%22%5D&frequency=M&type=value-Base&from=1993-01&to=2025-12>). After being transformed into logarithms, the variable was seasonally adjusted using additive Census X13 on the analyzed sample.
q	Volume of cigarettes taxed per capita. Volumes of cigarettes taxed with the combined tax rate, publicly available from the IFP MFSR Tax returns – tobacco products database (source: <https://www.mfsr.sk/en/finance/institute-financial-policy/economy-statistics/taxes-contributions/>), were divided by monthly population data, publicly available from the SOSR Datacube database (source: <http://datacube.statis-tics.sk/#!/view/en/vbd_dem/om7102mr/v_om7102mr_00_00_00_en>). After being transformed into logarithms, the variable was seasonally adjusted using additive Census X13 on the analyzed sample.
Wage	Real monthly wages in 2015 prices. Both nominal monthly average wages and monthly HICP deflators for all products are publicly available from the NBS Macroeconomic database (source: <https://nbs.sk/statisticke-udaje/vybrane-makroekonomicke-ukazovatele/makroekonomicka-databaza/?timeSeriesId=%5B%221925%22%5D&fre-quency=M&type=valueBase&from=1993-01&to=2025-12>). After being transformed into logarithms, the variable was seasonally adjusted using additive Census X13 on the analyzed sample.
Spec. Duty	Real specific part of the tax rate on cigarettes in 2015 prices. Nominal specific part of the tax on cigarettes is publicly available from the IFP MFSR Tax returns – tobacco products database (source: <https://www.mfsr.sk/en/finance/institute-financial-policy/economy-statis-tics/taxes-contributions/>), which was adjusted using monthly HICP deflators for all products, publicly available from the NBS Macroeconomic database (source: <https://nbs.sk/statisticke-udaje/vybrane-makroekonomicke-ukazovatele/mak-roekonomicka-databaza/?timeSeriesId=%5B%221925%22%5D&frequency=M&type=val-ueBase&from=1993-01&to=2025-12>). After being transformed into logarithms, the varia-ble was seasonally adjusted using additive Census X13 on the analyzed sample.
HTP Duty	The tax rate on heated tobacco products in 2015 prices. Nominal tax rate on heated tobacco products is publicly available from the IFP MFSR Tax returns – tobacco products database (source: <https://www.mfsr.sk/en/finance/institute-financial-policy/economy-statistics/taxes-contributions/), which was adjusted using monthly HICP deflators for all products, publicly available from the NBS Macroeconomic database (source: <https://nbs.sk/statisticke-udaje/vybrane-makroekonomicke-ukazovatele/makroekonomicka-databaza/?timeSeriesId=%5B%221925%22%5D&frequency=M&type=value-Base&from=1993-01&to=2025-12>). After being transformed into logarithms, the variable was seasonally adjusted using additive Census X13 on the analyzed sample.

Note: The analyzed sample spans from Jun 2011 to May 2023, i.e., 144 observations with a monthly frequency.

T a b l e A2

Stationarity and Unit Root Tests

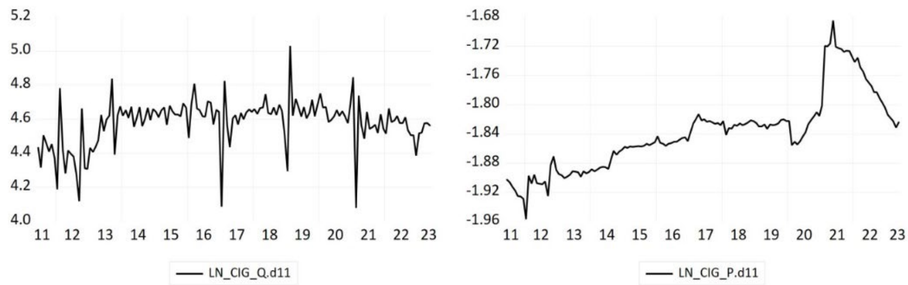
	KPSS	PP	ADF	DF-GLS
p	1.116**	–1.574	–1.635	–0.896
q	0.466**	–11.901**	–2.781	–1.534
Wage	1.331**	–0.926	–0.902	0.406
Spec. Duty	0.746**	–2.12	–2.004	–1.587
HTP Duty	1.195**	–0.926	–0.946	–0.188

Note: In Table A2, KPSS stands for Kwiatkowski-Phillips-Schmidt-Shin stationarity test with Newey-West bandwidth using Bartlett kernel for estimating residual spectrum, for which the Lagrange multiplier statistic is reported; PP stands for Phillips-Perron unit root test with Newey-West bandwidth using Bartlett kernel for estimating residual spectrum, for which the adjusted t-statistic is provided; ADF is Augmented Dickey-Fuller unit root test using Schwartz information criterion for selection of the number of lagged difference terms, regarding which conventional t-statistic is reported; and DF-GLS stands for the Elliott-Rothenberg-Stock variation of Dickey-Fuller generalized least squares unit root test using Schwartz information criterion for selection of the number of lagged difference terms, regarding which conventional t-statistic is provided. In all cases, only the constant is included among the exogenous terms, and all tests are conducted using the dependent variables at levels (without any differencing). Statistical significance at 5% level is indicated by **. Most of the results indicate that all variables are integrated of order one (stationary at first difference), with the exception of the PP for the variable *q*. The sample used for testing spans from Jun 2011 to May 2023, i.e., 144 observations with a monthly frequency.

Source: Author's own estimates based on the data specified in Table A1.

F i g u r e A1

Plot of the VECM Endogenous Variables for the Analyzed Sample



Note: The analyzed sample spans from Jun 2011 to May 2023, i.e., 144 observations with a monthly frequency.

Source: The sources of the data are specified in Table A1.