Tax-Induced Spatial Substitution and the Price Elasticity of Fuel Demand

7 July 2025

Ander Iraizoz, Oxford University Centre for Business Taxation

José M. Labeaga, Department of Economics, UNED

Working paper | 2025-03

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Ander Iraizoz José M. Labeaga

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Abstract

This paper estimates the price elasticity of fuel demand in the presence of spatial variations in fuel taxes. Exploiting variations in fuel taxes across Spanish regions, we obtain two main results. First, we document substantial spatial substitution in diesel sales, with an elasticity reaching -14 near regional borders. Second, we demonstrate that failing to account for such spatial substitution can lead to over-estimating the fuel price elasticity of demand. When spatial substitution is ignored, we estimate an elasticity of -3, while diesel demand becomes inelastic once spatial substitution is controlled for. Our findings highlight the relevance of spatial substitution on the observed fuel demand responses to unilateral tax changes, and suggest a more limited effectiveness of localized environmental policies aimed at reducing CO_2 emissions.

JEL classification: D12, H23, H71, H73, Q35, Q41.

Keywords: Fuel Taxation, Demand Elasticity, Spatial Substitution.

*Iraizoz: Centre for Business Taxation, Saïd Business School and Regent's Park College, University of Oxford, ander.iraizoz@sbs.ox.ac.uk. Labeaga: Department of Economics, UNED, jlabeaga@cee.uned.es. The results of this paper previously circulated as "Incidence and Avoidance Effects of Spatial Fuel Tax Differentials: Evidence using Regional Tax Variation in Spain". We thank Miguel Almunia, Antoine Bozio, Marika Cabral, Ron Chan, Courtney Coile, Michael Devereux, Massimo Filippini, Peter Haan, Julien Grenet, Niels Johannesen, Xavier Labandeira, Antoine Levy, Jordi Teixidó and participants at seminars and conferences at the University of Oxford, Paris School of Economics, Atlantic Workshop on Energy and Environmental Economics, CESifo Area Conference in Public Economics, Mannheim Conference on Energy and the Environment, Meeting on Public Economics and Annual Congress of the IIPF for their valuable comments and suggestions. The authors acknowledge funding from project PID2022-136376OB-I00. All remaining errors are our own.

1 Introduction

In light of growing climate concerns, it is essential to assess policy tools to reduce transport fuel use. Fuel taxes play a key role in internalizing CO_2-related emissions through higher pump prices. Their effectiveness depends on the price elasticity of fuel demand. Thus, accurate elasticity estimates are crucial for projecting environmental impacts and for designing efficient and cost-effective climate policies. However, identifying the causal effect of fuel prices on demand is empirically challenging. A common strategy exploits spatial tax variation, comparing regions experiencing tax changes to neighboring areas that do not (Li et al., 2014; Rivers and Schaufele, 2015; Tiezzi and Verde, 2016). Comparability of these regions improves if they share fuel supply costs or broader economic conditions, i.e., when they are part of an integrated market. Yet it is precisely in these integrated markets that spatial tax differences can induce cross-border substitution, as consumers shift purchases to lower-tax jurisdictions. This behavior can confound true demand responses to prices with spatial reallocation of sales, highlighting the need for strategies that ensure regional comparability and explicitly account for spatial substitution.

This paper estimates how fuel sales respond to spatial substitution incentives, and examines their impact on fuel price elasticity estimates based on spatial fuel tax variation. We exploit Spain's setting, where Spanish autonomous communities (regions from now on) adjusted regional fuel excise taxes from 2002 to 2019. We analyze 33 regional tax changes using various two-way fixed effects (TWFE) methods and monthly data on fuel prices and sales across Spanish provinces from 2007 to 2020. Our main analysis focuses on diesel since this accounts for 82% of total liters of automotive fuel sales in Spain[1].

We find that cross-regional price gaps lead to significant spatial substitution in diesel sales. The general elasticity of spatial substitution is -2.8, rising to -14.4 within 25 km of a regional border. This indicates that spatial substitution is strongest near borders, but extends beyond them. Our strategy offers two advantages over prior studies (Banfi et al., 2005; Leal et al., 2009; Jansen and Jonker, 2018; Teixidó et al., 2024). First, within-country tax variation enables more credible comparisons between treated and control units, favoring

[1] Historically, diesel excise taxes have been considerably lower than those on gasoline, leading to a predominance of diesel-powered vehicles in Spain. Diesel is also the primary fuel used in professional transportation.

the internal validity of our estimates. Second, estimating distance-dependent elasticities allows us to characterize spatial substitution patterns, enhancing the external validity of our results for other federal systems with different distributions of petrol stations.

This spatial substitution response has important implications for estimating the price elasticity of fuel demand. When spatial substitution is ignored, we estimate an elasticity of –3.0, but accounting for it reduces the estimate below unity. We further validate this finding using a restricted sample of non-border provinces where substitution is limited, for whom we estimate an elasticity of -0.7. To the best of our knowledge, we provide the first evidence highlighting the need to account for spatial substitution when using geographic tax variation, contributing to the literature on the estimation of fuel demand elasticities (Davis and Kilian, 2011; Coglianese et al., 2017; Levin et al., 2017; Knittel and Tanaka, 2021). Our findings also speak to the evidence that fuel demand is more responsive to tax-induced price changes than to equivalent supply-side shocks (Li et al., 2014; Rivers and Schaufele, 2015 Tiezzi and Verde, 2016). Since many studies rely on spatial tax variation, our results suggest substitution may partly explain this differential response[2].

Our findings have important implications for assessing the effectiveness of diesel taxes in reducing CO_2 emissions (Sterner, 2007; Davis and Kilian, 2011; Andersson, 2019). Ignoring spatial substitution suggests a 1% diesel price increase cuts emissions by 1.6 million tons annually in Spain. However, accounting for substitution reduces this estimate to just 0.4 million tons. This stark contrast underscores the importance of incorporating spatial substitution effects when assessing the climate impact of unilateral fuel tax policies.

Finally, our results speak to the importance of coordination in environmental policy design. Considering that fuel taxes face limited public support (Oates and Portney, 2003; Carattini et al., 2019; Douenne and Fabre, 2022), decentralized adoption may improve their acceptability (Oates, 1999; Agrawal et al., 2024). This logic underpinned the introduction of Spain's regional fuel tax band, whose revenues were earmarked for healthcare, earning its popular name the *health cent*. However, we show that unilateral regional fuel taxes can trigger cross-border substitution, undermining their effectiveness. This highlights the poten-

[2]Andersson (2019) is a notable exception, finding stronger responses to taxes in a setting less sensitive to spatial substitution.

tial cost of decentralizing climate policy, calling for interjurisdictional coordination to reduce leakage and improve the impact of environmental tax measures.

The paper proceeds as follows. Section 2 presents a two-market model illustrating price differential effects. Section 3 outlines Spain's fuel tax system. Section 4 describes the data. Section 5 introduces the empirical strategy. Section 6 reports the main findings. Section 7 concludes.

2 Effects of fuel taxes under spatial substitutes

We present a simple partial-equilibrium model to conceptualize the fuel demand responses to fuel taxes under the possibility of spatial substitution. We assume that geographical locations represent different markets that sell a homogeneous product, each geographic market being a substitute for one another. This framework is well suited to retail fuel markets, where geographic differentiation plays a central role (Slade, 1998), but the model could also be applied to other markets with imperfect substitutes[3]. More formal price competition models under geographical differentiation can be found in Anderson et al. (2001) or Bajo-Buenestado and Borrella-Mas (2019).

2.1 Model setup

We consider two markets, A and B, each located in different jurisdictions with distinct per-unit taxes τ_A and τ_B, respectively. Both markets sell a homogeneous product to consumers who can choose between the two suppliers. Let $c(Q_i)$ represent the common marginal cost of producing and distributing fuel, excluding taxes. The effective marginal cost for each firm includes the jurisdiction-specific tax $MC_i(Q_i) = c(Q_i) + \tau_i$ for $i = \{A, B\}$. Firms set their retail prices p_A and p_B simultaneously.

Market demand. We decompose the demand in market A into two components:

$$D_A(p_A, p_B, d_{AB}) = D_A(p_A|_{p_A=p_B}) + \Delta D_A(p_A - p_B, d_{AB}) \quad (1)$$

[3]Other forms of substitution could include fuel used for other purposes, as in Marion and Muehlegger (2008)

where $D_A(p_A)|_{p_A=p_B}$ represents the demand in market A without spatial substitution incentives and $\Delta D_A(p_A - p_B, d_{AB})$ represents the demand-shift between markets A and B, which depends on price differences $p_A - p_B$ and distance between markets d_{AB}. By assuming separability and linearity relative to price differences, the demand can be expressed as follows:

$$D_A(p_A, p_B, d_{AB}) = D_i(p_A|_{p_A=p_B}) + (p_A - p_B)f(d_{AB}) \tag{2}$$

where $f(d_{AB})$ represents the function of the effect of price differences depending on the distance between markets A and B.

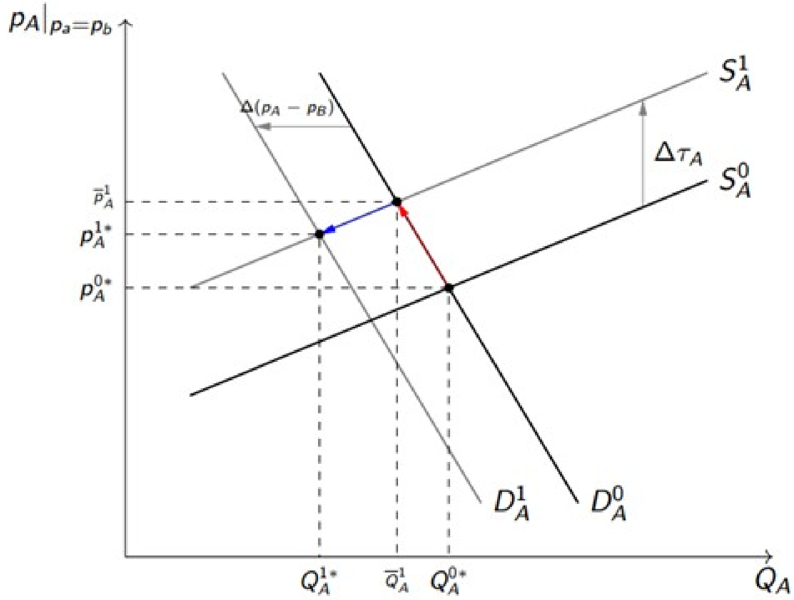
2.2 Effect of unilateral tax changes

Figure 1 illustrates the effect of unilateral tax changes on quantities, disentangling price and substitution effects. The y-axis plots the price under $p_A = p_B$, reflecting the demand curve when prices in all substitutable markets adjust simultaneously. An increase in market A's tax τ_A raises its marginal cost MC_A, shifting the supply curve upward. The resulting changes in quantities can be decomposed into the two key effects that we aim to identify empirically: the price effect and the spatial substitution effect.

Price effect. The tax change results in a movement along the demand curve $D_A(p_A, p_B)$, holding $p_A = p_B$ constant. This captures the direct response of fuel demand to changes in local fuel prices, abstracting from any spatial substitution.

Spatial substitution effect. A change in the tax differential between the two jurisdictions, $\tau_A - \tau_B$, leads to a change in the relative price, $p_A - p_B$. This shift in the price differential induces consumers to reallocate demand from market A to market B, causing a leftward shift in the demand curve $D(p_A, p_B)$. As a result, both the equilibrium price and quantity in market A decline. As suggested in Equation 2, the extent of the shift will depend on the costs of relocating demand, including distance, as well as other determinants of market integration (Bajo-Buenestado and Borrella-Mas, 2019).

Figure 1: Illustration of the effect of a fuel tax change applied to market A



Notes: The figure illustrates the effect of a tax change applied to market A on its prices and quantities. The red arrow represents the demand effect, indicating movements along the demand curve following a tax change. The blue arrows represent spatial substitution effect, capturing the demand shifts because of changes to spatial substitution incentives. The y-axis plots the price under $p_A = p_B$, reflecting the demand curve when prices in all substitutable markets adjust simultaneously.

3 Institutional design

This section outlines the key institutional features of this study. First, we provide a general overview of the fuel tax framework in Spain. Second, we describe the regional fuel tax band introduced in 2002, which generates the quasi-experimental variation exploited in this study.

3.1 Fuel taxation in Spain

Fuel taxation in Spain comprises value-added taxes (VAT) and fuel excise taxes, which are set both at the central and regional level. The effect of fuel taxes on retail petrol station prices can be expressed as follows:

$$P_{irt} = (P_{irt}^{pre} + \overline{\tau}_t + \tau_{rt}) \times (1 + VAT_t(\%)) \tag{3}$$

where P_{irt} represents the retail fuel price in petrol station i located in region r at time

t, P_{irt}^{pre} represents the pre-tax fuel price in petrol station i located in region r at time t, $\overline{\tau}_t$ represents the fuel excise tax set by the central government at time t, τ_{rt} represents the regional excise fuel tax applied to sales in region r at time t, and VAT_t represents the VAT rate applied to fuel at time t, which applies nationally[4].

Over our study period from 2007 to 2020, central government fuel excise taxes $\overline{\tau}_t$ were set at 38 cents per liter for diesel and 47 cents per liter for gasoline. The lower excise taxes for diesel than gasoline have prompted diesel to be more prevalent among vehicles in Spain[5]. In 2019, diesel represented 82.41% of the total liters of automotive fuel sales in Spain. Therefore, our analysis concentrates on the effects of diesel taxation.

3.2 Regional fuel excise tax band

In 2002, the Spanish Government introduced a regional fuel excise tax band[6], which allowed Spanish regions to levy a fuel excise tax of up to 4.8 cents per liter[7]. The effective burden of the tax was amplified by the VAT rate, implying that regions could influence fuel taxes by up to 5.808 cents per liter. The tax band became popularly known as the *health cent*, as revenues were earmarked for funding regional healthcare services.

Between 2007 and 2019, there were 33 changes to regional fuel excise taxes. Figure 2 illustrates the adoption of the tax band over time and across provinces. By 2012, nine regions, mostly in the south, had increased their regional diesel excise taxes. In 2012, several additional regions raised their rates, with eight reaching the maximum allowable rate by early 2013. Between 2013 and 2018, some regions decreased their regional excise taxes, while others such as Galicia and Aragón further increased them. Finally, in 2019, the regional tax band was mandated to the maximum of 4.8 cents per liter. This change affected eight regions, five of which were applying no regional fuel tax.

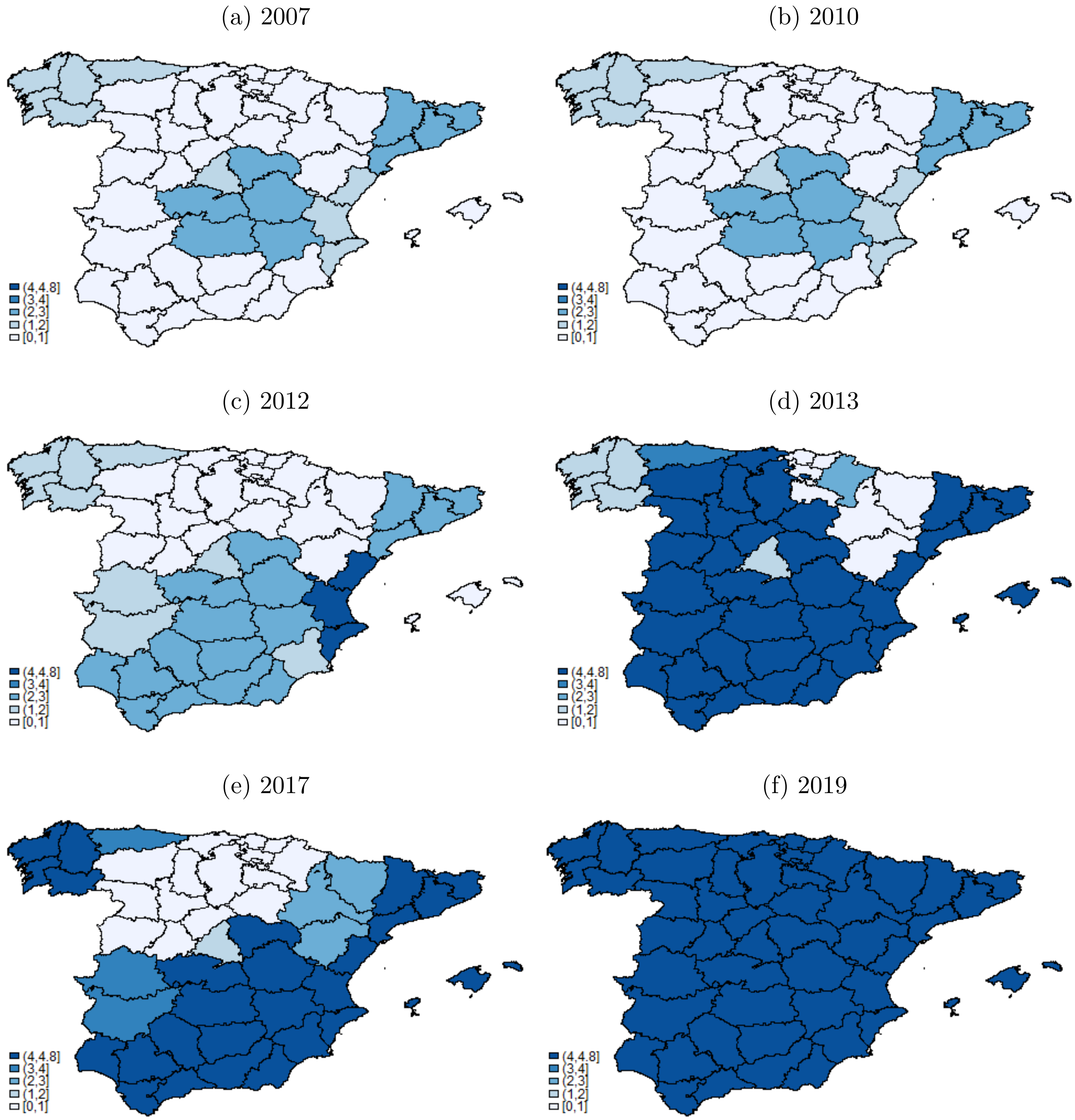
[4]The VAT rate was set at 16% in 2007-2010, to 18% in 2010-2012, and finally 21% after 2012.

[5]Overall, in 2021 fuel taxes represented 49% of retail diesel prices and 53% of gasoline prices in Spain, which is lower than the Eurozone average of 56% and 61%, respectively (Excise duties (europa.eu)).

[6]The legal name of the tax was the Tax on Retail Sales of Certain Mineral Oils (*Impuesto sobre las Ventas Minoristas de determinados Hidrocarburos*, IVMDH), and was passed by the Law 24/2001, of 27th of December, of Fiscal, Administrative and Social Order Policies.

[7]The ceiling for the regional excise fuel tax was extended over time. This was 1.7 cents per liter in 2002-2003, 2.4 cents per liter in 2004-2007 and 4.8 cents per liter since 2008.

Figure 2: Regional excise tax for automotive diesel between 2007-2019.



Notes: The figure displays the regional band of diesel excise taxes for Spanish provinces on the 1st January in 2007 (panel a), 2010 (panel b), 2012 (panel c), 2013 (panel d), 2017 (panel e) and 2019 (panel f). The automotive taxes for gasoline correspond almost identically.
Source: Spanish Ministry of Ecological Transition and Demographic Challenge.

4 Data

Our primary data comes from the Spanish National Markets and Competition Commission (*Comisión Nacional de Mercados y Competencia*, CNMC). The CNMC collects fuel prices and fuel sales aggregated at the province-level for every month since 2007. This dataset is based on information sent by petrol station brands operating in Spain to the Spanish Ministry of Ecological Transition and Demographic Challenge. The CNMC dataset includes average monthly fuel prices and total liters of sales by type of diesel and gasoline in each of the 52 Spanish provinces. Despite the aggregate nature of the data, the availability of fuel price and sales with such a time and geographic disaggregation remains an exception in Europe[8].

Our analysis requires constructing the fuel prices and taxes for each province's closest competitors in different regions. These are calculated using a weighting the fuel prices and taxes from adjacent provinces in different regions as follows:

$$\begin{aligned} \overline{P}_{jt} &= \sum_{k \neq i} w_{i_p k} P_{kt} \\ \overline{\tau}_{jt} &= \sum_{k \neq i} w_{i_p k} \tau_{kt} \end{aligned} \tag{4}$$

where $\overline{P}_{jt}$ and $\overline{\tau}_{jt}$ are, respectively, the diesel prices and taxes for the combination of the closest cross-regional competitor provinces k of province p, and $w_{i_p k}$ is the share of petrol station i in province p whose closest cross-regional competitor is in province k. To construct these variables, we use data on the location of all petrol stations in Spain coming from the *Geoportal*[9]. We identify the closest cross-regional competitor of each petrol station using Picard (2010). Additionally, we use Huber and Rust (2016) to calculate the distance by car between cross-regional petrol station pairs using OpenStreetMap data.

Sample selection. We exclude petrol stations in Ceuta, Melilla and the Canary Islands from our analysis since they have different indirect tax regimes, which limits their compara-

[8]To the best of our knowledge, Italy is the only other European country providing monthly fuel data at a similar level of disaggregation, but this is only available from 2015. While state-level data are available for the US, the data for Spain remains more granular as Spanish provinces are on average 20 times smaller than U.S. states (196,670 km^2/US state vs. 9,730.6 km^2/Spanish province).

[9]The *Geoportal* data contains the location of all petrol stations and can be accessed in real time (link).

bility with the rest of Spain. We also restrict our data to the period between January 2007, when both price and sales data became available, and February 2020, just before the onset of the Covid pandemic. This results in 7,584 province-month observations.

Summary statistics. Table B.1 presents the summary statistics for the variables used in our analysis. The relevant variables include diesel taxes, prices and sales from the CNMC, distance shares to closest cross-regional competitors calculated using *Geoportal* data, we well as additional control variables from the Spanish Office of National Statistics.

5 Empirical strategy

In this section, we outline the empirical strategies used to assess the effects of regional variation in diesel taxes in Spain. First, we introduce our two-stage least squares (2SLS) specification to estimate the elasticities of diesel sales with respect to diesel prices and spatial price differentials. Second, we present a linear panel event study framework to evaluate the dynamic effects of regional diesel tax changes on diesel prices and sales over time.

5.1 Fuel demand elasticities

We estimate a specification on the elasticity of fuel sales with respect to fuel prices and spatial fuel substitution incentives:

$$\ln S_{pt} = \alpha_p + \gamma_t + \varepsilon^p \ln P_{pt} + \varepsilon^s (\ln P_{pt} - \ln \overline{P}_{jt}) + \mathbf{x_{pt}} \theta + \varepsilon_{pt} \tag{5}$$

where $\ln S_{pt}$ denotes the log of diesel sales in province p and in month t, α_p and γ_t represent province and month fixed effects, respectively; $\ln P_{pt}$ is the log of diesel prices in province p and at time t; and $\overline{P}_{jt}$ captures the weighted average diesel prices of the closest cross-region competitors j of petrol stations in province p in month t. The vector $\mathbf{x_{pt}}$ includes time-varying province-level controls such as the employment rate, the logarithm of population, and the logarithm of real GDP per capita. Standard errors are clustered at the province level to account for within-province serial correlation.

Since diesel prices are endogenous to diesel sales, we implement a 2SLS estimation strategy, using regional variations in fuel taxes as instruments for fuel prices. Specifically, we instrument $\ln P_{pt}$ and $(\ln P_{pt} - \ln \overline{P}_{jt})$ with their respective fuel tax counterparts τ_{pt} and $(\tau_{pt} - \overline{\tau}_{jt})$. The first stage regression specifications corresponding to the 2SLS estimation are reported in Appendix C.1. The parameters of interest are ε^{p} and ε^{s}, which can be interpreted as follows. ε^{p} captures the price elasticity of fuel demand, holding constant spatial substitution incentives. This coefficient reflects the relevant parameter for assessing the effectiveness of fuel taxes in reducing fuel consumption. In contrast, ε^{s} measures the spatial substitution elasticity of fuel sales in response to cross-regional price differentials, holding constant their own price.

Distance ranges. To assess how the distance between petrol stations along regional borders influences the effects of spatial substitution, we re-estimate Equation 5 by allowing the elasticity of spatial fuel substitution to vary across distance bands to the nearest cross-regional petrol station.

$$\begin{aligned} \ln S_{pt} =& \alpha_p + \gamma_t + \varepsilon^{p} \ln P_{pt} + \varepsilon^{s,\underline{d}} (\ln P_{pt} - \ln \overline{P}_{jt}) \times \pi_{p}^{\underline{d}} + \\ & + \varepsilon^{s,\overline{d}} (\ln P_{pt} - \ln \overline{P}_{jt}) \times (1 - \pi_{p}^{\underline{d}}) + \mathbf{x_{pt}} \theta + \varepsilon_{pt} \end{aligned} \tag{6}$$

where $\pi_{p}^{\underline{d}} = Pr(d_{i_p} \leq d)$ denotes the share of petrol stations i in province p whose nearest cross-regional competitor is located within a distance of d km. To examine how the responsiveness to spatial substitution incentives varies with distance, we estimate a series of regressions using alternative distance thresholds. The parameters $\varepsilon^{s,\underline{d}}$ and $\varepsilon^{s,\overline{d}}$ can be interpreted as distance-specific elasticities of spatial substitution. Specifically, $\varepsilon^{s,\underline{d}}$ captures the elasticity of spatial substitution incentives for petrol stations located within d km of their nearest cross-regional competitor, holding constant their $\ln(P_{jt})$. Conversely, $\varepsilon^{s,\overline{d}}$ measures the elasticity of spatial substitution incentives for petrol stations located beyond d km threshold, also holding their $\ln(P_{jt})$ constant. This distinction allows us to assess how the strength of spatial substitution incentives varies with proximity to regional borders.

5.2 Event study regressions

We estimate linear panel event study regressions to provide graphical evidence on the exogeneity of regional excise tax changes with respect to pre-existing time trends in diesel prices and sales in Spain. Our event study framework allows us to assess the dynamic response of outcomes around tax changes and to test for the presence of differential pretrends. The implementation of our event study approach follows Freyaldenhoven et al. (2021):

$$y_{it} = \alpha_i + \gamma_t + \sum_{m=-G}^{M} \beta_m \tau_{i,t-m} + \mathbf{x_{it}}\theta + \varepsilon_{it} \tag{7}$$

where y_{it} denotes the outcome variable of interest, namely diesel prices, $y_{it} = P_{it}$, and the log of diesel sales, $y_{it} = \ln S_{it}$, for observational unit i and time t[10]. The specification includes fixed unit effects α_i and fixed time effects γ_t, and control for unobserved heterogeneity between units and time. $\tau_{i,t-m}$ stands for regional diesel taxes m periods prior to time t for unit i. The vector $\mathbf{x_{it}}$ includes control variables relevant to each dataset[11]. We normalize coefficients relative to period $m = 1$ when estimating fuel price responses, and to $m = 2$ when estimating fuel sales responses, addressing the possibility of anticipation to the reform (Coglianese et al., 2017).

6 Results

We present our results on the effects of diesel taxes on diesel sales under spatial substitution incentives. First, we estimate the dynamic effects of diesel excise taxes on diesel prices and sales in an event study framework, which allows us to assess the validity of our identification strategy. Second, we present our baseline 2SLS estimates of the elasticity of fuel demand with respect to fuel prices and spatial substitution incentives. Third, we explore the heterogeneity of the estimated price elasticity of diesel demand for subsamples differing in the possibility of spatial substitution. Fourth, we show the implications of our fuel price elasticity estimates

[10] For the *Geoportal* dataset, the observational unit i corresponds to individual petrol stations, with t measured in days. For the CNMC dataset, i refers to Spanish provinces, and t corresponds to months.

[11] For the *Geoportal* data, this includes indicators of competitive intensity faced by each petrol station. For CNMC data, it includes province-level employment rates, the logarithm of population and the logarithm of real GDP per capita.

on CO_2 emissions and environmental governance. Finally, we perform a series of robustness checks to examine the sensitivity of our findings to alternative model specifications and sample restrictions.

6.1 Event study evidence

The fundamental assumption underlying the use of regional diesel taxes as exogenous variations in diesel prices is that the timing of these changes are orthogonal to pre-existing regional trends in diesel prices and sales. To assess the plausibility of this assumption, Figure 3 displays event study estimates of the dynamic effect of regional diesel tax changes on diesel prices and diesel sales.

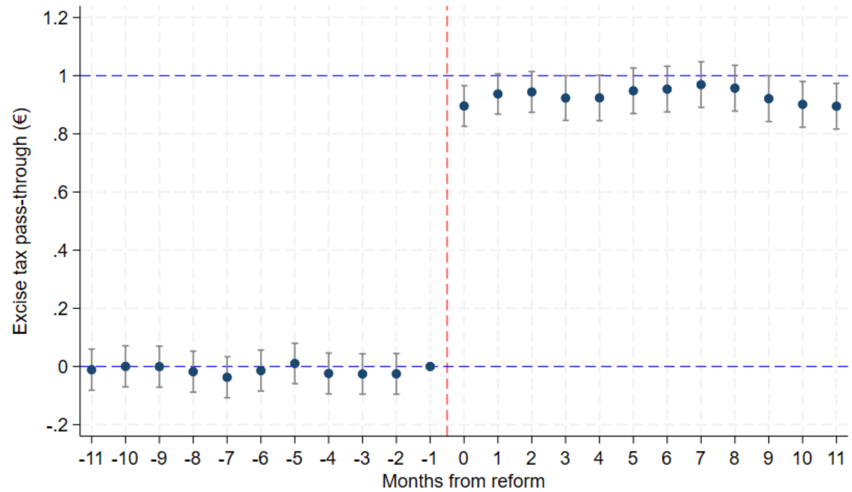
Pass-through. Figure 3a displays the plot of the event study on the dynamic effect of diesel taxes on diesel prices. First, we did not observe differential regional diesel price trends before the implementation of diesel excise tax changes. Second, we document a rapid and near-complete pass-through of diesel taxes to diesel prices. Within 12 months of a tax change, between 90% and 95% of the tax is reflected in diesel prices. These results align with prior studies on fuel tax pass-through (Chouinard and Perloff, 2004; Alm et al., 2009; Marion and Muehlegger, 2011).

Sales response. Figure 3b illustrates the event study estimates of the dynamic effects of changes in the diesel tax on diesel sales. First, we find no evidence of differential trends in diesel sales across regions before tax changes, which supports the assumption that regional diesel tax reforms can be treated as exogenous shocks. Second, we detect anticipatory behavior by consumers, who increase diesel purchases in the month preceding the changes in diesel taxes[12] (Coglianese et al., 2017). Third, we estimate that a one percentage point increase in diesel excise taxes leads to a relative decline in diesel sales of approximately 2% in regions experiencing a tax change relative to those that do not. The response appears both immediate and persistent following the reform.

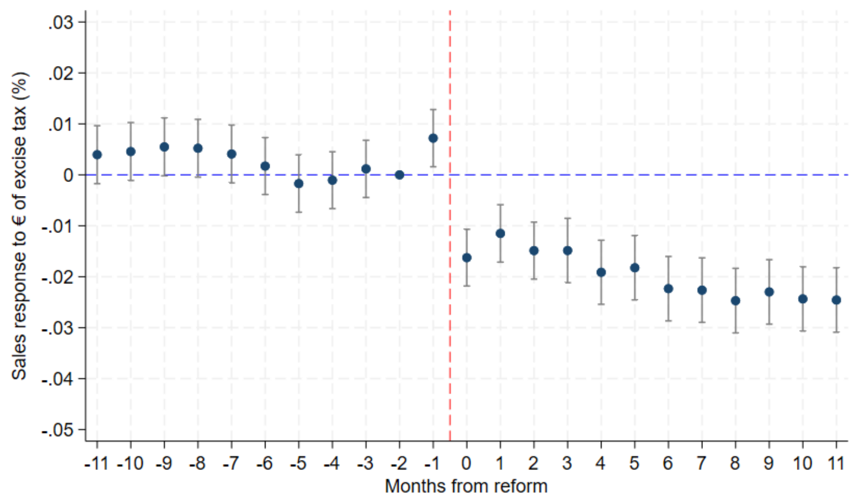
[12]To account for this behavior, our TWFE regressions exclude observations from the month before and after the excise tax changes.

Figure 3: Event study evidence on price and sales responses to regional diesel excise taxes

(a) Pass-through



(b) Diesel sales



Notes: The figure shows the event study graph on the dynamic effect of the diesel excise taxes on diesel prices using (panel a), as well as the dynamic effects of diesel taxes on sales responses (panel b). The vertical dashed red line refers to the period of the reform.
Source: Spanish National Markets and Competition Commission (CNMC).

6.2 Price elasticities and spatial substitution

In this subsection, we describe our results on the elasticity of diesel sales with respect to diesel prices and spatial substitution incentives. Table 1 reports these 2SLS results for a number of specifications. First of all, we verify that the instrument relevance condition is satisfied across all specifications, and reject that the instruments are weak. While the Stock-

Yogo critical values from Stock and Yogo (2005) are only available for cases with one or two endogenous regressors, they are known to decrease with the number of endogenous variables. Therefore, the available critical value for two endogenous regressors provides a conservative upper bound. We compare our Kleibergen and Paap (2006) Wald F-statistics in columns 3-6 to the Stock-Yogo 10% maximal IV size threshold of 7.03. In all cases, our F-statistics exceed this benchmark, suggesting that our instruments are sufficiently strong.

Table 1: 2SLS results on diesel price and substitution elasticities

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-3.05***	-0.43	-0.51	-0.49	-0.35	-0.34
	(0.63)	(1.48)	(1.74)	(1.65)	(1.58)	(1.54)
$\ln P_{pt} - \ln \overline{P}_{jt}$		-2.80*				
		(1.22)				
$(\ln P_{pt} - \ln \overline{P}_{jt})\times$ $\Pr(d_{i_p} \leq d)$			-14.35***	-6.96**	-4.79**	-3.93**
			(3.87)	(2.37)	(1.63)	(1.43)
$(\ln P_{pt} - \ln \overline{P}_{jt})\times$ $\Pr(d_{i_p} > d)$			-1.32	-1.28	-1.32	-1.37
			(1.31)	(1.33)	(1.29)	(1.28)
First stage (F-Stat)	445.82	18.75	11.85	12.41	12.70	12.70
Endogenous variables	1	2	3	3	3	3
N (obs)	7,382	7,382	7,382	7,382	7,382	7,382

Notes: This table provides 2SLS results on the elasticity of diesel sales with respect to diesel prices and diesel price differentials to neighbouring provinces. Column (1) does not control for spatial substitution incentives, while column (2) does. Columns (3)-(6) weight the effects of price differentials by the share petrol stations within distance ranges to the closest cross-regional competitors. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

In column 1, we estimate a diesel price elasticity without controls for spatial substitution incentives, which provides an estimate of -3.05. This is substantially larger than typical estimates reported in the literature (Sterner, 2006; Brons et al., 2008; Dahl, 2012; Labandeira et al., 2017). However, when we account for spatial substitution incentives (column 2), the estimated price elasticity falls below unity and is no longer statistically significant at conventional levels. Instead, we find that most of the response is driven by the price differential relative to cross-regional competitors. The estimated elasticity of diesel sales with respect to spatial substitution incentives is -2.80. These findings indicate that failing to account for cross-regional substitution may lead to a substantial overestimation of the price elasticity of diesel demand. Our results underscore the importance of controlling for spatial arbitrage opportunities to obtain more accurate estimates of fuel price responsiveness.

Columns 3 to 6 report the results of the elasticity of spatial substitution incentives disaggregated by distance to the nearest cross-regional competitor. This disaggregation facilitates extrapolation to settings with different cross-border spatial configurations than those observed across Spanish provinces. We estimate an elasticity of diesel cross-regional substitution incentives as high as -14.4 for petrol stations located within 25km of a cross-regional border. As the distance increases, the estimated elasticities decline substantially, but remain economically and statistically significant at conventional levels. For petrol stations within 100 km of a border, the elasticity is estimated at -3.9. Beyond this range, the elasticity remains above unity but is no longer statistically distinguishable from zero. These results suggest that spatial substitution responses are highly concentrated near regional borders, yet their aggregate effects extend well beyond immediate border areas.

Our estimated elasticities of spatial substitution incentives are notably larger than those reported in previous studies. For instance, Banfi et al. (2005) estimate an elasticity of -1.5 for gasoline sales along Swiss borders, while Jansen and Jonker (2018) find only modest substitution effects across Dutch international borders. The greater magnitude of our estimates may reflect the higher degree of market integration within countries relative to international settings, where cultural and logistical frictions may still be more pronounced. As economic integration in the EU continues to strengthen, our results may serve as a benchmark for the potential effects of cross-border tax differentials. In addition, our data include diesel sales used for professional transportation, which is more responsive to tax differentials due to the high mobility of heavy-duty vehicles[13] [14].

[13] Banfi et al. (2005) focused on gasoline sales, which is predominantly used by private vehicles, whereas diesel powers a significant share of trucks and buses. Teixidó et al. (2024) provide evidence of strategic fuel purchases by heavy-duty vehicles in response to a diesel tax increase in Portugal, which affected diesel sales in Spain. Similarly, Jansen and Jonker (2018) rely on household-level consumption data, which excludes professional transportation.

[14] Allocating diesel taxation based on consumption rather than point of sale could address the spatial mobility of professional fleets, although evidence from the US suggests implementation challenges may persist (Marion and Muehlegger, 2018).

Gasoline. We also study the effects of spatial substitution on gasoline sales, which are detailed in Appendix E. The results shows that while also relevant, the effects of spatial substitution are substantially larger for diesel. The difference in results could be driven by the fact that diesel is mainly used for transportation and long-distance travel, being substantially more sensitive to spatial substitution incentives.

6.3 Price elasticities on subsets of provinces

We provide additional evidence on the impact of cross-regional substitution possibilities on the estimation of fuel price elasticities when these are identified using spatial variation in fuel taxes. We estimate the following simple 2SLS regression with TWFE:

$$\ln S_{pt} = \alpha_p + \gamma_t + \varepsilon^p \ln P_{pt} + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{8}$$

where we again instrument for $\ln P_{pt}$ using regional variations in τ_{it}. We estimate this regression across various subsets of provinces that differ in their potential for cross-regional substitution. First, we focus on three economically integrated areas, including (i) the *central* area integrated around Madrid, (ii) *northern* area, which gives access to the Atlantic coast, and (iii) *eastern* area, which connects to the Mediterranean coast. Second, we estimate elasticities using data from provinces that do not share borders with other regions, which implies very limited possibilities of cross-regional spatial substitution.

Table 2 reports the 2SLS estimates of diesel price elasticities for these different subsets of provinces. In columns 2-4, we find that the estimated price elasticities exceed 3 in these integrated areas, particularly in the Atlantic and Mediterranean regions, where freight transport by trucks is more prevalent because of access to seaports. By contrast, column 5 shows that the estimated elasticity in provinces with limited opportunities of spatial substitution is substantially lower, at -0.72. This figure lies within the range commonly reported in the literature (Sterner, 2006; Brons et al., 2008; Dahl, 2012; Labandeira et al., 2017).

Our findings relate to prior studies documenting a larger behavioral response to tax-induced price changes than to equivalent supply-side shocks (Li et al., 2014; Rivers and Schaufele, 2015; Tiezzi and Verde, 2016). These studies exploit spatial variation in fuel

Table 2: 2SLS price elasticities for different subsets of provinces

	All Spain	Integrated areas			No border
		Center	Atlantic	Mediterranean	
	(1)	(2)	(3)	(4)	(5)
$\ln P_{pt}$	-3.05***	-3.48***	-5.69***	-4.58*	-0.72*
	(0.63)	(0.93)	(1.51)	(1.99)	(0.31)
First stage (F-Stat)	445.82	199.35	19.27	253.69	100.19
Endogenous variables	1	1	1	1	1
N (obs)	7,382	1,074	1,554	1,226	1,238

Notes: This table provides the diesel price elasticities for different subsets of provinces in Spain, including all provinces in Spain apart from the Canary Islands, Ceuta and Melilla (column 1), central provinces (Madrid, Toledo, Avila, Segovia, Guadalajara, Cuenca and Soria) (column 2), the area around the Atlantic coast (Cantabria, Basque Country, Navarre, Rioja, Burgos and Huesca) (column 3), Mediterranean coast (Aragon, Catalonia and Valencian Community) (column 4) and provinces not sharing a border with another region (Balearic Islands, Girona, Barcelona, A Coruña, Pontevedra, Valladolid, Cadiz and Malaga) (column 5). The Kleibergen-Paap rk Wald F-statistic is reported for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

taxes and attribute the differential response to the greater salience of taxes. However, cross-border tax differences may themselves be especially salient to drivers, making them more likely to adjust their refueling behavior by purchasing fuel in lower-tax areas, rather than by substantially reducing overall fuel consumption. Our results suggest that spatial substitution may partly explain the apparent overreaction to tax-induced price changes relative to supply-side shocks.

6.4 Fuel price elasticities and CO_2 emissions

Our results show significant cross-regional substitution, which affects the elasticities estimated from spatial tax variation. We investigate the implications on the measurement of the effect of fuel prices on CO_2 emissions and on environmental governance.

CO2 emissions. We evaluate how the estimated price elasticities affect the estimates on the effectiveness of fuel tax policies in reducing CO_2 emissions. We illustrate the environmental relevance of this issue by comparing CO_2 reductions under two scenarios. Using 2019 diesel consumption (2.32×10^{10} liters), diesel's carbon intensity (2.24 kg CO_2 per liter; EIA, 2023), and elasticities from Table 2, a 1% price increase yields a 1.6 Mt CO_2 cut using a naive elasticity (–3.05 of column 1), but only 0.37 Mt using an adjusted elasticity (–0.72, column 5). Thus, ignoring substitution overstates reductions fourfold. This finding highlights the importance of incorporating spatial behavioral responses into the evaluation of carbon pricing policies, particularly in settings with high internal market integration.

Environmental governance. Our findings also have implications for environmental governance. Fuel taxes and other forms of environmental taxation often face limited public support (Oates and Portney, 2003; Carattini et al., 2019; Douenne and Fabre, 2022). In response, the literature on fiscal federalism has argued that decentralized adoption may enhance political acceptability (Oates, 1999; Agrawal et al., 2024). This rationale underpinned the introduction of Spain's regional fuel tax band, with revenues earmarked for healthcare—earning it the popular name *health cent*. We show, however, that unilateral regional fuel taxes can induce cross-regional substitution, undermining their effectiveness. This underscores a potential cost of decentralizing climate policy and highlights the need for interjurisdictional coordination to limit leakage and enhance the effectiveness of environmental taxation.

The issue of environmental governance is particularly salient in the EU, as the Effort Sharing Regulation (ESR) comes into effect. The ESR sets binding national emission reduction targets for each EU member state[15], placing national governments under pressure to implement effective climate policies. In countries like Spain, where national-level accountability coexisted with regional tax autonomy, accurate elasticity estimates would have been crucial for assessing mitigation potential and minimizing policy leakage. More broadly, at the European level, cross-border substitution is likely to undermine the effectiveness of environmental taxation, contributing to carbon leakage and reducing the overall impact of climate policy.

6.5 Robustness checks

In this subsection, we conduct several robustness checks for our results.

TWFE methods. We implement the TWFE estimator proposed by De Chaisemartin and d'Haultfoeuille (2020), which addresses potential issues arising from treatment effect

[15] The ESR collectively aims for a 30% reduction in emissions by 2030 relative to 2005 levels.

heterogeneity across units and over time. As shown in Figure 3, the sharp and stable responses to fuel tax changes suggest that treatment effect heterogeneity is likely limited in our context. Consistently, Table D.1 shows that our baseline TWFE results are virtually unchanged when applying this alternative estimator.

Petrol station weights. We assess the sensitivity of our results to the use of province-level petrol station weights in the regressions, which assign greater influence to provinces with a higher number of petrol stations. Tables D.2 and D.3 confirm that our main findings and interpretations remain robust under this weighting scheme.

Anticipation. In our main analysis, we exclude the observations immediately before and after excise tax changes to address anticipation effects (Coglianese et al., 2017). Tables D.4 and D.5 show that our results remain largely unchanged when these observations are retained. Furthermore, Tables D.6 and D.7 show that the results are also robust when explicitly controlling for the elasticity of intertemporal substitution in fuel sales.

Separate effects of $\ln(P_{it})$ and $\ln(\overline{P}_{jt})$. Consistent with our model in Equation 2, we estimate the effects of own prices and cross-regional price differentials. As a sanity check, we estimate the regression by separately including $\ln(P_{it})$ and $\ln(\overline{P}_{jt})$. Table D.8 confirms that the estimated coefficients are virtually identical to those obtained in our main specification. This check validates our approach of modeling spatial substitution incentives through the price differential, as the separate effects are consistent with the interpretation of a substitution margin driven by relative price levels.

7 Conclusion

This paper examined spatial substitution in fuel sales and its impact on estimating fuel price elasticity. Exploiting regional diesel tax variation in Spain, we used TWFE methods to jointly estimate fuel price elasticity and cross-regional substitution. Our results provide new evidence on how regional tax differentials shape fuel demand behavior. We find that spatial substitution significantly shifts the geographic distribution of diesel sales, with elasticity exceeding 14 for stations within 25 km of regional borders. Fuel price elasticities are highly sensitive to these incentives: ignoring substitution yields an elasticity of –3.0, while accounting for it lowers estimates below unity. These findings highlight the risk of overstating price responsiveness when cross-regional substitution is not properly considered.

Our findings have key implications for diesel taxation as a tool to reduce CO_2 emissions. First, from a measurement standpoint, ignoring spatial substitution can overstate the effectiveness of fuel taxes. Second, regionally implemented policies may be less effective, as cross-border purchases can offset local tax increases. This concern is particularly relevant in the EU, where increasing economic integration and country-specific climate goals may lead to a mismatch between fiscal instruments and the spatial distribution of emissions. Evidence from regional variation in Spain illustrates how such dynamics can play out.

Several questions remain open. Future research should explore the mechanisms driving spatial fuel substitution, such as fuel tourism, commuting, and professional transport. Further empirical work is also needed to quantify how tax-driven substitution affects the effectiveness of fuel taxes in reducing emissions.

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Appendix to

Tax-Induced Spatial Substitution and the Price Elasticity of Fuel Demand

Ander Iraizoz and José M. Labeaga

July 7, 2025

The appendix provides additional institutional details (Appendix A), additional details on the summary statistics (Appendix B), additional details on the empirical strategy (Appendix C), additional results on sales responses (Appendix D), and additional results on the price elasticities for gasoline (Appendix E).

A Institutional design: Regional fuel taxes

This appendix provides additional details on the fuel taxation system in Spain. In 2002, the Spanish Government introduced the Tax on Retail Sales of Certain Mineral Oils[16] (*Impuesto sobre las Ventas Minoristas de determinados Hidrocarburos*, IVMDH). This regional excise tax band allowed Spanish regions to set a fuel excise tax of in excess of the central government's fuel excise tax. From 2002 to 2003, the ceiling for the regional excise fuel tax was 1.7 cents per liter, from 2004 to 2007 the ceiling was 2.4 cents per liter and since 2008 this is set at 4.8 cents per liter. Sixteen regions had competence to modify this tax, while the Canary Islands, Ceuta and Melilla did not since they have a separate indirect tax regime.

In 2013, the IVMDH was integrated into the general Excise Duty on Mineral Oils (*Impuesto Especial sobre Hidrocarburos*, IEH) to comply with European Union (EU) Law. The IVMDH was ruled unconstitutional as it was justified based on budgetary purposes, while EU Law dictates that fuel excise taxes should aim to influence fuel consumption based, for instance, on environmental objectives. In 2014, the European Court of Justice ruled that the IVMDH was unconstitutional and requested the return of its revenues between 2002 and 2013.

[16] Ley 24/2001, de 27 de diciembre, de Medidas Fiscales, Administrativas y del Orden Social.

B Additional details: Summary statistics

Table B.1: Summary statistics on main variables using CNMC data

	Mean (1)	S.D. (2)	Min (3)	Max (4)	N(obs) (5)
Diesel products					
Retail price (c/litre)	116.28	15.16	82.18	146.57	7,584
Pre-tax price (c/litre)	55.68	10.85	30.16	77.87	7,584
Tax component (c/litre)	60.60	7.04	46.37	71.30	7,584
Regional excise tax (c/litre)	2.98	2.55	0.00	5.81	7,584
Tax differential (c/litre)	0.43	2.23	-5.81	5.81	7,584
Price differential (c/litre)	0.20	2.22	-8.02	7.87	7,584
Total sales (1000 litres)	38,259.92	34,284.29	3,955.61	218,952.81	7,584
Distance range to competitors (%)					
$Pr(d_{i_p} \leq 25km)$	13.94	18.81	0.00	87.50	7,584
$Pr(d_{i_p} \leq 50km)$	31.24	30.59	0.00	100.00	7,584
$Pr(d_{i_p} \leq 75km)$	47.30	36.32	0.00	100.00	7,584
$Pr(d_{i_p} \leq 100km)$	61.66	37.23	0.00	100.00	7,584
Additional variables					
Petrol stations	183.56	135.43	37.00	704.00	7,584
Population	920,187.59	1,168,045.79	89,415.00	6,747,068.00	7,584
Real GDP (2016€)	22,710.36	4,927.66	14,525.44	40,748.40	7,584
Employment rate (%)	47.17	5.53	32.78	63.12	7,584

Notes: The table describes the summary statistics for the main variables of the CNMC dataset. The relevant variables include diesel price, taxes and sales from the CNMC, distance shares to closest cross-regional competitors calculated using *Geoportal* data, we well as additional control variables from the Spanish Office of National Statistics. The data covers the period between January 2007 and February 2020. This considers monthly data for all Spanish provinces, excluding the Canary Islands, Ceuta and Melilla.
Source: Spanish National Markets and Competition Commission (CNMC).

C Additional details: Empirical strategy

This appendix provides additional details on the empirical strategy for estimating sales responses to cross-regional differences in fuel prices.

C.1 First-stage regression for 2SLS approach

We describe the first stage of the 2SLS regression to estimate the price elasticity of diesel sales. We present the two specifications that we present.

Specification 1: Aggregate effects of spatial substitution incentives. We provide the first-stage specification for our estimation of Equation 5, where we have two endogenous price variables and two tax instruments. The first stage regressions are specified as follows:

$$\ln P_{pt} = \alpha_p + \gamma_t + \beta_1 \tau_{pt} + \beta_2(\tau_{pt} - \overline{\tau}_{jt}) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{C.1}$$

$$\ln P_{pt} - \ln \overline{P}_{jt} = \alpha_p + \gamma_t + \beta_3 \tau_{pt} + \beta_4(\tau_{pt} - \overline{\tau}_{jt}) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{C.2}$$

where $\ln P_{pt}$ represents the log of diesel prices in province p and in month t, $\overline{P}_{jt}$ represents diesel prices of the closest cross-regional competitor province j and in month t, which is a weighted average of province diesel prices of the closest cross-regional competitors of province p. α_p represents province fixed effects, γ_t represents month fixed effects, τ_{pt} represents the diesel tax rate in province p and in month t, $\overline{\tau}_{jt}$ represents the diesel tax rate of the closest cross-regional competitor province j and in month t. $\mathbf{x_{pt}}$ includes province level employment rate, the logarithm of population and the logarithm of real GDP per capita. We use clustered standard errors at the province level.

Specification 2: Distance weighted effects. We provide the first-stage specification for our estimation of Equation 6, where we have three endogenous price variables and three tax instruments. The first stage regressions are specified as follows:

$$\ln P_{pt} = \alpha_p + \gamma_t + \beta_1 \tau_{pt} + \beta_2(\tau_{pt} - \overline{\tau}_{jt}) \times \pi_p^d + \beta_3(\tau_{pt} - \overline{\tau}_{jt}) \times (1 - \pi_p^d) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{C.3}$$

$$(\ln P_{pt} - \ln P_{jt}) \times \pi_p^d = \alpha_p + \gamma_t + \beta_1 \tau_{pt} + \beta_2(\tau_{pt} - \overline{\tau}_{jt}) \times \pi_p^d + \beta_3(\tau_{pt} - \overline{\tau}_{jt}) \times (1 - \pi_p^d) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{C.4}$$

$$(\ln P_{pt} - \ln P_{jt}) \times (1 - \pi_p^d) = \alpha_p + \gamma_t + \beta_1 \tau_{pt} + \beta_2(\tau_{pt} - \overline{\tau}_{jt}) \times \pi_p^d + \beta_3(\tau_{pt} - \overline{\tau}_{jt}) \times (1 - \pi_p^d) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{C.5}$$

where $\pi_p^d = Pr(d_{i_p} \leq d)$ stands for the share of petrol stations i in province p whose closest cross-regional competitor is closer than d km. The rest of the variables have already been described.

D Additional results: Robustness checks

This appendix provides robustness checks on sales responses using alternative specifications.

D.1 TWFE estimator

We compare our baseline two-way fixed effects (TWFE) estimator to the estimators developed by the emerging TWFE literature. Recent studies have shown that heterogeneity in treatment effects over time and across units could bias ordinary TWFE estimators. In particular, the literature is concerned about cases when already treated units enter the control group, as increasing dynamic effects would contaminate the control group. In our setting, the event study evidence provided in Figure 3, which shows immediate and stable effects over time, suggests heterogeneity is likely to play a limited role.

Despite important developments in the literature, most studies are only applicable to binary treatments and staggered adoption (Sun and Abraham, 2021; Borusyak et al., 2024; Callaway and Sant'Anna, 2021). However, our fuel tax treatments can be multiple per treatment unit and treatment is continuous. Therefore, we check our estimations with the estimator developed by De Chaisemartin and d'Haultfoeuille (2020), which is compatible with multiple and continuous treatments. We estimate the following regression simple TWFE regression to check the robustness of our estimation:

$$y_{pt} = \alpha_p + \gamma_t + \beta\tau_{pt} + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{D.6}$$

where y_{pt} represents the outcomes of interest, which are diesel prices, $y_{pt} = P_{pt}$, and the log of diesel sales, $y_{pt} = \ln S_{pt}$, in province p and month t. α_p refers to province fixed effects and γ_t refer to monthly fixed effects. τ_{pt} refer to regional diesel taxes in province p and month t. $\mathbf{x_{it}}$ includes province level employment rate, the logarithm of population and the logarithm of real GDP per capita. We use clustered standard errors at the province level.

Table D.1 shows that our simple TWFE estimator provides almost identical point estimates as the estimator developed by De Chaisemartin and d'Haultfoeuille (2020), while estimation is more precise in our simple TWFE estimator. This gives confidence on the validity of our baseline TWFE estimator for the rest of our analysis.

Table D.1: Comparison of the TWFE estimators

	Baseline TWFE method (1)	De Chaisemartin & D'Hautefeuille (2020) (2)
P_{pt}	0.853***	0.829***
	(0.038)	(0.162)
$\ln S_{pt}$	−0.018***	-0.018***
	(0.003)	(0.005)

Notes: The table provides the TWFE estimation of effect of diesel taxes on diesel prices and on the logarithm of diesel sales by province for our baseline TWFE estimator (column 1) and for the TWFE estimator developed by De Chaisemartin and d'Haultfoeuille (2020) (column 2). Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

D.2 Petrol station weights

We assess the sensitivity of our results to using province level weights for the total petrol stations from whom the fuel data is collected. Table D.2 and Table D.3 show that the estimated effects are slightly smaller, but the results and their interpretation remain very similar to our baseline results.

Table D.2: 2SLS results on diesel price and substitution elasticities, using petrol station weights

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-2.46***	-0.54	-0.31	-0.42	-0.36	-0.43
	(0.42)	(1.03)	(1.23)	(1.13)	(1.14)	(1.09)
$\ln P_{pt} - \ln P_{jt}$		-2.17*				
		(0.98)				
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} \leq d)$			-11.86**	-4.78*	-4.09**	-3.39**
			(4.19)	(1.92)	(1.57)	(1.31)
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} > d)$			-1.17	-1.23	-0.95	-0.80
			(0.95)	(0.91)	(0.91)	(0.89)
First stage (F-Stat)	343.47	14.12	7.50	8.90	8.72	9.26
Endogenous variables	1	2	3	3	3	3
N (obs)	7,382	7,382	7,382	7,382	7,382	7,382

Notes: This table provides 2SLS results on the elasticity of diesel sales with respect to diesel prices and diesel price differentials to neighbouring provinces by the share petrol stations within distance ranges to the closest cross-regional competitors. Province-level observations are weighted by the number of petrol stations. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

Table D.3: Weighted 2SLS price elasticities for different subsets of provinces, using petrol station weights

	All Spain	Integrated			No border
		Center	North	East	
	(1)	(2)	(3)	(4)	(5)
$\ln P_{pt}$	-2.46***	-4.88***	-4.42***	-4.34*	-0.67
	(0.42)	(0.60)	(0.97)	(1.83)	(0.45)
First stage (F-Stat)	343.47	493.82	36.35	233.27	55.25
Endogenous variables	1	1	1	1	1
N (obs)	7,382	1,074	1,554	1,226	1,238

Notes: This table provides the diesel price elasticities for different subsets of provinces in Spain, including all provinces in Spain apart from the Canary Islands, Ceuta and Melilla (column 1), central provinces (Madrid, Toledo, Avila, Segovia, Guadalajara, Cuenca and Soria) (column 2), the northern area (Cantabria, Basque Country, Navarre, Rioja, Burgos (Castile-Leon) and Huesca (Aragon)) (column 3), eastern area (Aragon, Catalonia and Valencian Community) (column 4) and provinces not sharing a border with another region (Balearic Islands, Girona, Barcelona, A Coruña, Pontevedra, Valladolid, Cadiz and Malaga) (column 5). Province-level observations are weighted by the number of petrol stations. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

D.3 Anticipation

We also assess the sensitivity of results when not dropping the observations immediately before and after an excise tax change, which are dropped in our baseline specification to control for anticipation. Table D.4 and Table D.5 show that dropping these observations to account for the effect of anticipation to the reform has a very small effect on the estimated price elasticities.

Table D.4: 2SLS results on diesel price and substitution elasticities, including anticipation observations

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-3.03***	-0.36	-0.43	-0.42	-0.28	-0.27
	(0.59)	(1.45)	(1.69)	(1.61)	(1.54)	(1.50)
$\ln P_{pt} - \ln P_{jt}$		-2.85*				
		(1.21)				
$(\ln P_{pt} - \ln P_{jt})\times$ $\Pr(d_{i_p} < d)$			-14.01***	-6.83**	-4.74**	-3.93**
			(3.73)	(2.30)	(1.59)	(1.41)
$(\ln P_{pt} - \ln P_{jt})\times$ $\Pr(d_{i_p} < d)$			-1.44	-1.41	-1.45	-1.49
			(1.29)	(1.31)	(1.28)	(1.26)
First stage (F-Stat)	494.78	19.11	12.13	12.69	12.95	12.92
Endogenous variables	1	2	3	3	3	3
N (obs)	7,584	7,584	7,584	7,584	7,584	7,584

Notes: This table provides 2SLS results on the elasticity of diesel sales with respect to diesel prices and diesel price differentials to neighbouring provinces by the share petrol stations within distance ranges to the closest cross-regional competitors. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

Table D.5: 2SLS price elasticities for different subsets of provinces, including anticipation observations

	All Spain	Integrated			No border
		Center	North	East	
	(1)	(2)	(3)	(4)	(5)
$\ln P_{pt}$	-3.03***	-3.48***	-5.56***	-4.52*	-0.76*
	(0.59)	(0.85)	(1.46)	(1.98)	(0.31)
First stage (F-Stat)	494.78	245.64	20.61	248.64	113.38
Endogenous variables	1	1	1	1	1
N (obs)	7,584	1,106	1,580	1,264	1,264

Notes: This table provides the diesel price elasticities for different subsets of provinces in Spain, including all provinces in Spain apart from the Canary Islands, Ceuta and Melilla (column 1), central provinces (Madrid, Toledo, Avila, Segovia, Guadalajara, Cuenca and Soria) (column 2), the northern area (Cantabria, Basque Country, Navarre, Rioja, Burgos (Castile-Leon) and Huesca (Aragon)) (column 3), eastern area (Aragon, Catalonia and Valencian Community) (column 4) and provinces not sharing a border with another region (Balearic Islands, Girona, Barcelona, A Coruña, Pontevedra, Valladolid, Cadiz and Malaga) (column 5). Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

Controlling for anticipation. We further estimate the effect of anticipation by including a control for anticipation $\ln P_{pt+1} - \ln P_{pt}$. Therefore, Equation 5 stays as follows:

$$\ln S_{pt} = \alpha_p + \gamma_t + \varepsilon^p \ln P_{pt} + \varepsilon^s(\ln P_{pt} - \ln \overline{P}_{jt}) + \varepsilon^t(\ln P_{pt+1} - \ln P_{pt}) + \mathbf{x_{pt}}\theta + \varepsilon_{pt} \tag{D.7}$$

where we instrument $\ln P_{pt+1} - \ln P_{pt}$ using anticipation tax variations $\tau_{pt+1} - \tau_{pt}$. ε^t measures the anticipation elasticity of fuel sales. The rest of the variables have already been described.

Tables D.6 and D.7 show that we estimate an elasticity of intertemporal substitution of diesel sales of 1.26, which is significant even at the 0.1% significance level. However, we can observe that the results on the price elasticity of fuel demand and the spatial substitution elasticity of fuel sales are hardly affected by the inclusion of the control for anticipation elasticities.

Table D.6: 2SLS results on diesel price and substitution elasticities, controlling anticipation

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-3.01***	-0.37	-0.44	-0.43	-0.29	-0.28
	(0.61)	(1.47)	(1.72)	(1.64)	(1.57)	(1.53)
$\ln P_{pt+1} - \ln P_{pt}$	1.26***	1.08***	1.21***	1.18***	1.13***	1.13***
	(0.32)	(0.29)	(0.25)	(0.26)	(0.27)	(0.27)
$\ln P_{pt} - \ln P_{jt}$		-2.81*				
		(1.22)				
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} \leq d)$			-14.00***	-6.83**	-4.71**	-3.89**
			(3.68)	(2.29)	(1.60)	(1.42)
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} > d)$			-1.39	-1.35	-1.41	-1.45
			(1.31)	(1.34)	(1.30)	(1.28)
First stage (F-Stat)	2,225.24	12.46	8.95	9.32	9.49	9.47
Endogenous variables	2	3	4	4	4	4
N (obs)	7,488	7,488	7,488	7,488	7,488	7,488

Notes: This table provides 2SLS results on the elasticity of diesel sales with respect to diesel prices and diesel price differentials to neighbouring provinces by the share petrol stations within distance ranges to the closest cross-regional competitors. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

Table D.7: 2SLS price elasticities for different subsets of provinces, controlling anticipation

	All Spain	Integrated			No border
		Center	North	East	
	(1)	(2)	(3)	(4)	(5)
$\ln P_{pt}$	-3.01***	-3.44***	-5.55***	-4.46*	-0.71*
	(0.61)	(0.88)	(1.47)	(2.00)	(0.30)
$\ln P_{pt+1} - \ln P_{pt}$	1.26***	2.03***	0.44	2.27*	1.02
	(0.32)	(0.48)	(0.99)	(1.08)	(0.60)
First stage (F-Stat)	2,225.24	1,309.60	274.26	114.35	241.49
Endogenous variables	2	2	2	2	2
N (obs)	7,488	1,092	1,560	1,248	1,248

Notes: This table provides the diesel price elasticities for different subsets of provinces in Spain, including all provinces in Spain apart from the Canary Islands, Ceuta and Melilla (column 1), central provinces (Madrid, Toledo, Avila, Segovia, Guadalajara, Cuenca and Soria) (column 2), the northern area (Cantabria, Basque Country, Navarre, Rioja, Burgos (Castile-Leon) and Huesca (Aragon)) (column 3), eastern area (Aragon, Catalonia and Valencian Community) (column 4) and provinces not sharing a border with another region (Balearic Islands, Girona, Barcelona, A Coruña, Pontevedra, Valladolid, Cadiz and Malaga) (column 5). Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

D.4 Separate effects

We estimate the effects of fuel prices by separating the effects of own and competitor's taxes. Therefore, we estimate the following regression:

$$\begin{aligned}\ln S_{pt} =& \alpha_p + \gamma_t + \varepsilon^{o,\underline{d}}(\ln P_{pt} \times \pi_p^{\underline{d}}) + \varepsilon^{o,\overline{d}}(\ln P_{pt} \times (1 - \pi_p^{\underline{d}})) + \\ & + \varepsilon^{c,\underline{d}}(\ln \overline{P}_{jt} \times \pi_p^{\underline{d}}) + \varepsilon^{c,\overline{d}}(\ln \overline{P}_{jt} \times (1 - \pi_p^{\underline{d}})) + \mathbf{x_{pt}}\theta + \varepsilon_{pt}\end{aligned} \tag{D.8}$$

where $\varepsilon^{o,d}$ is the own price elasticity of fuel demand in distance d, keeping the competitor's price constant, i.e. allowing for price differentials. $\varepsilon^{c,d}$ is the elasticity of competitor's prices in distance d, keeping own prices constant.

The difference in empirical strategy is that Equation 6 estimates three parameters, while Equation D.8 estimates 4. Based on the coefficients in Equation D.8, we can construct the equivalence between the estimates:

$$\varepsilon^{p,d} = \varepsilon^{o,d} - \varepsilon^{c,d}$$

$$\varepsilon^{s,d} = -\varepsilon^{c,d}$$

Table D.8 shows that the estimated effects are equivalent to our main results in table 1. We can observe that column 2 offers the exact same estimates, while the results in columns 3-6 are quantitatively and qualitatively similar applying the equivalence of the estimates shown above.

Table D.8: 2SLS results on the elasticities of own and competitor's diesel prices

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-3.05***	-3.23***				
	(0.63)	(0.65)				
$\ln \overline{P}_{jt}$		2.80*				
		(1.22)				
$\ln P_{pt} \times Pr(d_{i_p} \leq d)$			-16.33**	-7.20**	-5.05***	-4.30***
			(5.86)	(2.19)	(1.11)	(0.90)
$\ln P_{pt} \times Pr(d_{i_p} > d)$			-1.82**	-1.77**	-1.53*	-1.81*
			(0.64)	(0.66)	(0.76)	(0.80)
$\ln \overline{P}_{jt} \times Pr(d_{i_p} \leq d)$			15.24**	6.90**	5.12**	3.73**
			(5.37)	(2.51)	(1.70)	(1.39)
$\ln \overline{P}_{jt} \times Pr(d_{i_p} > d)$			1.09	1.37	1.42	1.36
			(1.27)	(1.24)	(1.25)	(1.29)
First stage (F-Stat)	445.82	18.75	5.62	16.31	8.90	10.68
Endogenous variables	1	2	4	4	4	4
N (obs)	7,382	7,382	7,382	7,382	7,382	7,382

Notes: This table provides 2SLS results on the elasticity of diesel sales with respect to diesel prices and diesel price differentials to neighbouring provinces by the share petrol stations within distance ranges to the closest cross-regional competitors. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

E Additional results: Gasoline demand

This appendix describes the results on price and substitution elasticities for gasoline. We first present the event study evidence on the dynamic effects of gasoline taxes on gasoline prices and gasoline sales. Second, we describe our results on price and spatial substitution elasticities of gasoline demand.

E.1 Gasoline vs. diesel use in Spain

There are important differences in the use of gasoline- and diesel-powered vehicles in Spain. Historically, gasoline has been taxed at higher rates than diesel, which has led to a greater prevalence of diesel vehicles. By 2019, gasoline accounted for only 17.59% of total automotive fuel sales by volume. Diesel is the primary fuel used in professional transportation and is more commonly associated with long-distance travel. In contrast, gasoline tends to be used in high-performance and sports vehicles. As a result, spatial substitution incentives are likely to be more relevant for diesel than for gasoline.

E.2 Event study evidence

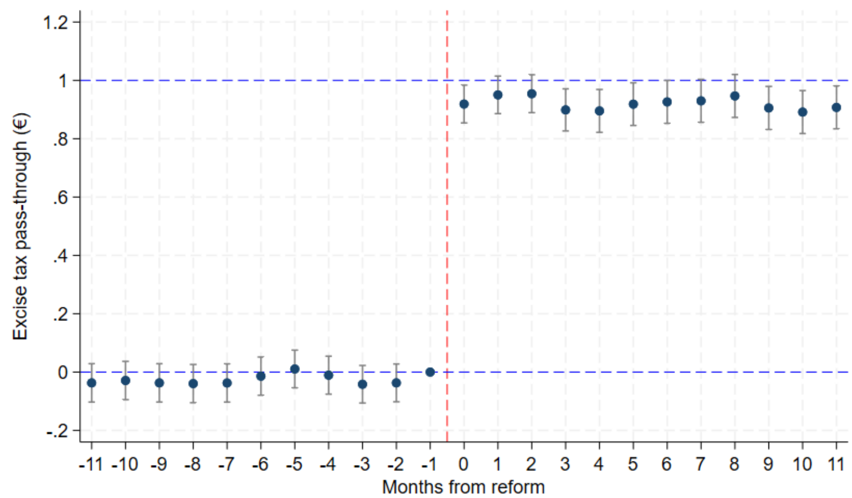
Figure E.1 displays the event study results on the dynamic effect of regional gasoline taxes on gasoline prices and sales. As with diesel, there is no evidence of differential regional trends prior to the gasoline tax reforms, either in prices or sales. Following the reform, we observe a near-complete and immediate pass-through of excise taxes to gasoline prices, with estimated pass-through rates ranging from 90% to 95%. These estimates closely mirror the pass-through observed for diesel.

We also document an immediate and stable decline in gasoline sales in treated regions relative to control regions. In addition, there is evidence of anticipatory behavior, with a significant sales response in the month preceding the tax change. While the overall pattern of responses is similar to that of diesel, the magnitude is smaller, suggesting that cross-regional tax differentials may induce weaker behavioral responses for gasoline than for diesel.

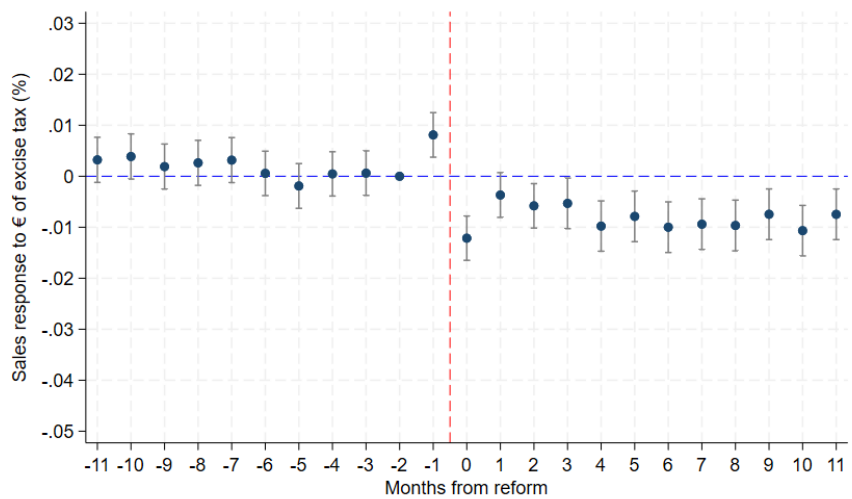
Overall, the absence of pre-trends in both prices and sales supports the identifying assumption that the timing of gasoline tax changes is plausibly exogenous to underlying regional trends.

Figure E.1: Event study evidence on price and sales responses to regional gasoline taxes

(a) Pass-through



(b) Gasoline sales



Notes: The figure shows the event study graph on the dynamic effect of the gasoline excise taxes on gasoline prices (panel a), as well as the dynamic effects of fuel taxes on gasoline sales responses (panel b). The vertical dashed red line refers to the period prior to the reform.
Source: Spanish National Markets and Competition Commission (CNMC).

E.3 Price elasticities of gasoline demand

Table E.1 reports our estimates on price and substitution elasticities of gasoline demand. Overall, the effects of spatial substitution incentives appear weaker for gasoline than for diesel. In column (1), we estimate a price elasticity of gasoline demand of -1.55 when spatial substitution is not accounted for. While this estimate is notably higher than typically reported in the literature, it remains approximately half the corresponding elasticity for diesel. Once we control for spatial substitution incentives in column (2), the estimated elasticity declines to below one and loses statistical significance at conventional levels. Notably, the effect of the price differential relative to cross-regional competitors is also statistically insignificant.

Columns (3) to (6) present a decomposition of the spatial substitution effect by distance to the nearest border. As with diesel, the point estimates are larger at shorter distances, but the magnitudes are less than half those for diesel and remain statistically insignificant at the 5% level. These findings suggest that spatial substitution plays a smaller role in gasoline markets. This is consistent with expectations in the Spanish context, where diesel fuels the vast majority of long-distance and transportation vehicles.

Table E.1: 2SLS results on sales response to spatial substitution incentives: Gasoline

			Distance range			
			$d = 25$	$d = 50$	$d = 75$	$d = 100$
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{pt}$	-1.55***	-0.87	-1.13	-1.12	-1.00	-0.92
	(0.37)	(1.21)	(1.28)	(1.25)	(1.23)	(1.24)
$\ln P_{pt} - \ln P_{jt}$		-0.70				
		(1.01)				
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} \leq d)$			-6.63	-3.01	-1.75	-1.36
			(3.96)	(2.10)	(1.45)	(1.27)
$(\ln P_{pt} - \ln P_{jt}) \times Pr(d_{i_p} > d)$			0.25	0.39	0.35	0.37
			(0.92)	(0.89)	(0.83)	(0.77)
First stage (F-Stat)	603.01	10.22	5.78	6.77	7.40	7.41
Endogenous variables	1	2	3	3	3	3
N (obs)	7,382	7,382	7,382	7,382	7,382	7,382

Notes: This table provides 2SLS results on the elasticity of gasoline sales with respect to gasoline prices and gasoline price differentials to neighbouring provinces by the share petrol stations within distance ranges to the closest cross-regional competitors. We report the Kleibergen-Paap rk Wald F-statistic for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).

E.4 Sensitivity with respect to spatial substitution

Table E.2 presents 2SLS estimates of the price elasticity of gasoline demand across subsets of provinces that differ in their exposure to spatial substitution opportunities. As with diesel, estimated elasticities are larger in more integrated areas with greater opportunity for spatial substitution. In contrast, gasoline demand appears inelastic in provinces that do not share a border with another autonomous community.

Table E.2: 2SLS price elasticities for different subsets of provinces: Gasoline

	All Spain	Integrated areas			No border
		Center	Atlantic	Mediterranean	
	(1)	(2)	(3)	(4)	(5)
$\ln P_{pt}$	-1.55***	-1.99***	-1.34	-1.72*	-0.14
	(0.37)	(0.60)	(0.76)	(0.76)	(0.33)
First stage (F-Stat)	603.01	351.32	29.48	347.09	410.80
Endogenous variables	1	1	1	1	1
N (obs)	7,382	1,074	1,554	1,226	1,238

Notes: This table provides the gasoline price elasticities for different subsets of provinces in Spain, including all provinces in Spain apart from the Canary Islands, Ceuta and Melilla (column 1), central provinces (Madrid, Toledo, Avila, Segovia, Guadalajara, Cuenca and Soria) (column 2), the area around the Atlantic coast (Cantabria, Basque Country, Navarre, Rioja, Burgos (Castile-Leon) and Huesca (Aragon)) (column 3), Mediterranean coast (Aragon, Catalonia and Valencian Community) (column 4) and provinces not sharing a border with another region (Balearic Islands, Girona, Barcelona, A Coruña, Pontevedra, Valladolid, Cadiz and Malaga) (column 5). The Kleibergen-Paap rk Wald F-statistic is reported for instrument relevance. Standard errors clustered at the province level in parenthesis. * p<0.05, ** p<0.01, *** p<0.001.
Source: Spanish National Markets and Competition Commission (CNMC).