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THE EFFECTS OF COMPETITION IN THE RETAIL GASOLINE INDUSTRY

Reid B. Taylor
Erich Muehlegger

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ABSTRACT

We estimate the effect of competition on incumbent firm pricing by using high frequency price data and the precise geographic location for all gas stations in California. Using an event study design, we find that the entry of a new station is associated with a 2.5 cent decrease in prices at incumbent stores, which equates to a 7% reduction in estimated retail markups. The effects are immediate, persistent, and show no sign of deterrence or limit pricing behavior. In contrast, nearby exit results in precisely estimated null effects on prices with no evidence of predatory pricing in the lead up to the station departure. The results are consistent across all fuel blends and dissipate with station distance. Finally, we explore the asymmetric effects, showing that the difference cannot be attributed to difference in branding, proximity to highway, or data quality idiosyncrasies, although we find suggestive evidence that exit tends to happen in more competitive markets and amongst less heavily trafficked stations.

Reid B. Taylor
Federal Reserve Bank of Dallas
reidbtaylor@gmail.com

Erich Muehlegger
University of California,Davis
Economics Department
One Shields Ave
Davis, CA 95616
and NBER
emuehlegger@ucdavis.edu

1 Introduction

The competitive effect of entry and exit is a central question in industrial organization. Although most models of competition predict that increased competition tends to lower prices, the impact of changes in competition depends on the nature of competition, as does the symmetry or asymmetry between the effects of entry and exit. In this paper, we estimate the effects of station entry and exit in local markets for retail gasoline in California. Quantifying the extent of these effects is of direct policy importance, given immediate concerns about high gasoline prices and competition within the state and longer-run interest in understanding how a shift towards alternative fuel vehicles might impact retail gasoline prices. Yet, theory on the relationship between market composition, entry, and prices in retail fuel markets provides ambiguous guidance ((Barron et al. 2004)). Moreover, empirically estimating the impact of changes in market size in this setting has traditionally been a challenge due to the endogeneity of market structure. Profit maximizing firms are attracted to markets with higher prices and profit margins. Moreover, entering and exiting firms might differ on both observable and unobservable characteristics. Endogenous selection biases cross-sectional estimates for the causal effect of market size on price ((Tappata and Yan 2017)).

To calculate the reduced-form causal effect of market size, we use a panel of daily station-level prices and the precise geographic location for the universe of California gas stations from 2014-2018 with geographic and temporal variation in exposure to changes in the number of nearby competitors through entry and exit. The resulting data set includes over 700 new station entry and station exit events and 35 million price observations. In the spirit of Arcidiacono et al. (2020), we use both difference-in-differences and event-study designs to compare the change in prices at incumbent stations before and after a competing station enters or exits the market, using stations that do not face entry or exit as the control group. We are able to control for the endogenous location decision of entering and exiting firms by including a rich panel of station and day-of-sample fixed effects and, city-specific linear time trends. Under the assumption that the exact timing of the entry or exit events is conditionally exogenous, the coefficient is identified by within-station variation in the number of nearby competitors. Our event study results support this key identifying assumption – markets with and without entry or exit follow parallel trends in the periods leading up to the respective event. In addition, we find little evidence of turnover, that the effect of entry or exit on competition is attenuated by entry or exit of other stations in the periods preceding, following, or simultaneous with the original entry and exit.

We find that increased market competition reduces prices, but that the effects of entry and exit are asymmetric. Entry of a new gas station nearby is associated with a 2.5 cent reduction in gas prices at nearby (less than 1 mile away) incumbent stations, representing a 7% reduction in average retail markups over the sample period. Our event study specification highlights that the effect of entry is immediate, occurring the month following the entry event, and persistent, lasting for years. It is present for all grades of gasoline and attenuates with the distance between the incumbent stations and the entrant, consistent with prior literature on the tight spatial nature of retail gasoline competition. Lastly, we find suggestive evidence of heterogeneity across types of stations – notably, that high volume hypermart stations (i.e. Costco) have a stronger entry effect on incumbents than branded and unbranded entrants.

In sharp contrast, we estimate precise null effects of station exit on nearby incumbent station pricing. Across grades, distances and station types, we find little evidence that exit events lead incumbent stations to raise prices. We explore this asymmetry by comparing the nature of entry and exit events. We find little evidence that the asymmetry in the effect of entry and exit is attributable to observable station characteristics or to the identification of entry and exit events. But we do find evidence that entry events tend to occur in more concentrated markets, where we estimate the impact of a change in competition to be greater. In addition, we find suggestive evidence that the exiting stations differ from incumbent (or entering) stations on unobservable dimensions, as reflected by the frequency of missing pricing observations pre-exit. If, prior to exit, these low-reporting stations tend to exert less competitive pressure on neighboring firms, their exit might not impact prices.

Our work contributes to a recent, growing literature that uses station-level data to estimate the effect of market size on retail fuel prices. Although there has long been interest in entry and exit in retail fuel markets, (e.g., (Barron et al. 2004), Tappata and Yan (2017)), estimation, in the spirit of Arcidiacono et al. (2020), has until recently been limited by a lack of granular price data in settings with changes in market structure ((Haucap et al. 2017)). Here, our work complements a growing series of papers that estimate the effect of station entry in Spain (Bernardo (2018), González and Moral (2023)), Germany ((Fischer et al. 2023)), Mexico ((Davis et al. 2023)) and Australia (Ormosi et al. (2024)) using high-frequency administrative data. We contribute to this literature in two respects.

First, recent papers focus almost exclusively on the impact of station entry on prices. Our paper provides some of the first evidence of the effect of both entry and *exit* events and is the first (to our knowledge) to highlight the asymmetry in the impacts of station entry and exit on incumbent prices.[1] Our focus on the asymmetric response to entry and exit complements the recent work by Ormosi et al. (2024), who study entry and exit effects in Western Australia, but focus on how the magnitude of entry and exit events relates to the income level of the local market.

The asymmetry in the impact of entry and exit is particularly relevant in light of the broader energy transition away from fossil fuels. As society moves towards reducing its reliance on fossil fuels, quantifying how markets will be impacted by a shift towards electrified transportation is paramount. Although gasoline demand is expected to decrease over time as the market penetration of electric vehicles increases, millions of gasoline-powered cars will still be operating on California's roads in the years to come. Our paper offers some preliminary insight into the potential short-term effects of market structure on the gasoline prices should the energy transition leads to falling demand for liquid transportation fuels and modest station exit. If, as our findings suggest, exiting stations have little impact on incumbent pricing, the short-term impacts of station exit might be muted.

[1] Although we are not aware of other papers that focus upon the asymmetric effects of entry and exit in retail gasoline, the asymmetric speed with which gasoline prices rise and fall, termed rockets and feathers, has long been studied. (see e.g., Borenstein et al. (1997), Lewis and Noel (2011))

In a related vein, an increasing number of local jurisdictions are seeking to impose supply-side restrictions as part of the energy transition. Starting with Petaluma in 2021, a number of local jurisdictions in California have banned the construction of new gas stations and restricted the expansion of existing stations. Although we find limited impacts of station exit, our results suggest that banning *entry* of new stations would lead to counterfactually higher prices.

Second, our work provides the first event-study-based causal estimates of the effect of market size and station entry on pricing in California. As such, our work contributes to the understanding of competition and market power in the California gasoline market. For the past two decades, gasoline in California has been significantly more expensive than gasoline in the rest of the country, after accounting for higher excise taxes, state-specific environmental regulations[2], and high entry costs due to zoning laws and high land values. This gap has drawn scrutiny from academics (e.g., (Borenstein et al. 2004) examines wholesale market power, (Hastings 2004; Taylor et al. 2010) examine vertical mergers, among others) and spurred investigations from the California Energy Commission[3] and the California Department of Justice[4]. In particular, the California Energy Commission report concludes that over the last several decades "the primary cause of the residual price increase is simply that California's retail gasoline outlets are charging higher prices" Although we find that incumbents lower their prices after the entry of a new competitor, the effect sizes are modest relative to the residual price gap estimated by the CEC.

The paper proceeds as follows: Section 2 presents the data used in the empirical study, Section 3 discusses the empirical strategy, Section 4 presents the estimation results, Section 5 discusses robustness and extensions, and Section 6 concludes.

[2]For three decades, stations in California, by regulation, must sell a special blend of gasoline (CARBOB), distinct from the fuel sold in surrounding states. California's unique gasoline blend regulations are served almost exclusively by the few in-state refineries.

[3]https://www.energy.ca.gov/sites/default/files/2019-11/Gas_Price_Report.pdf

[4]https://oag.ca.gov/antitrust/gasoline

2 Data

To calculate the effect of entry and exit on pricing, we analyze a panel of daily retail gasoline prices in California from the Oil Price Information Service (OPIS) covering the years 2014-2018. We observe daily prices for each fuel blend with the precise geographic coordinates for each station and a unique site identifier linked to each station's geographic location. The identifier remains consistent when a store undergoes renovations, changes ownership, or changes store or fuel branding, allowing us to not confound station changes that do not alter the number of competitors in the market with entry or exit. For convenience, in the rest of the paper we use the terminology "station" to refer to the unique site where a gas station is located.

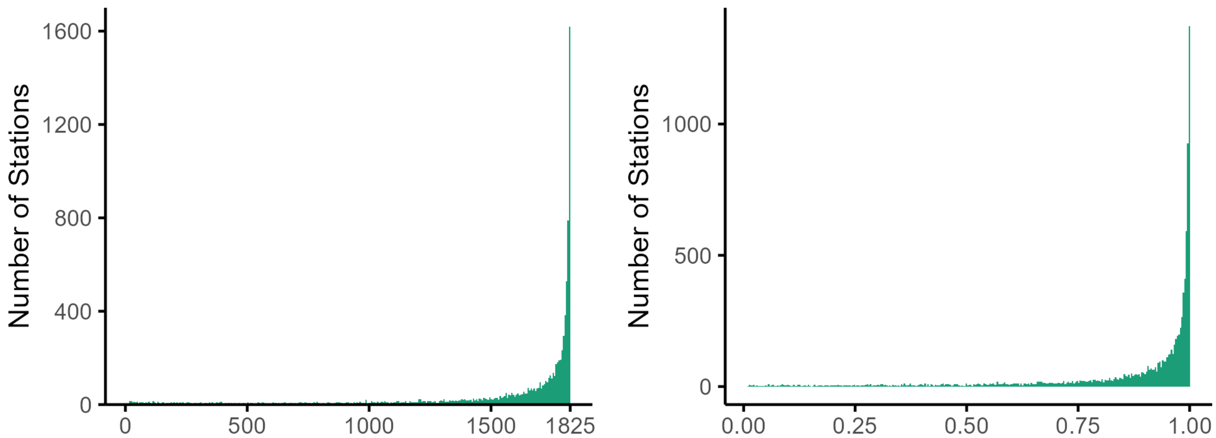
The data include information on 9,716 stations in California. We exclude stations with less than 14 total price observations over the 4-year sample period as well as four stations for which geographic coordinates are unavailable. This results in a panel of 9,539 stations, with over 35 million price observations across the three major grades of gasoline.

The OPIS data also reports the brand of fuel sold at each station. We follow the California Energy Commission in classifying brands into three categories.[5] ”Branded” stations sell gasoline under the brand of a company that refines petroleum products, specifically Chevron, Shell, 76, Exxon Mobil, and Valero. Branded fuels are marketed to consumers using refining company signage and include proprietary additives that are blended into the fuel.[6] The term ”unbranded” refers to gasoline sales that are sold under brands other than that of the refining company. Unbranded stations purchase gasoline from suppliers at the ”rack” (i.e., wholesale terminal) and are not contractually obligated to purchase branded gasoline from a specific refiner. Finally, we use the term ”hypermart” to refer to gasoline sold by Costco, Walmart, Sam's Club, Safeway or other large retailers.

[5] https://www.energy.ca.gov/sites/default/files/2020-02/2020-01_Petroleum_Watch.pdf

[6] Although we observe the brand under which gasoline is sold, we do not observe the whether the store is independently owned, franchised, directly owned by a petroleum company or sells gasoline under the brand on a consignment basis.

Figure 1: Price Observations by Station: 2014-2018



(a) Observation Count (b) % of Potential Observation

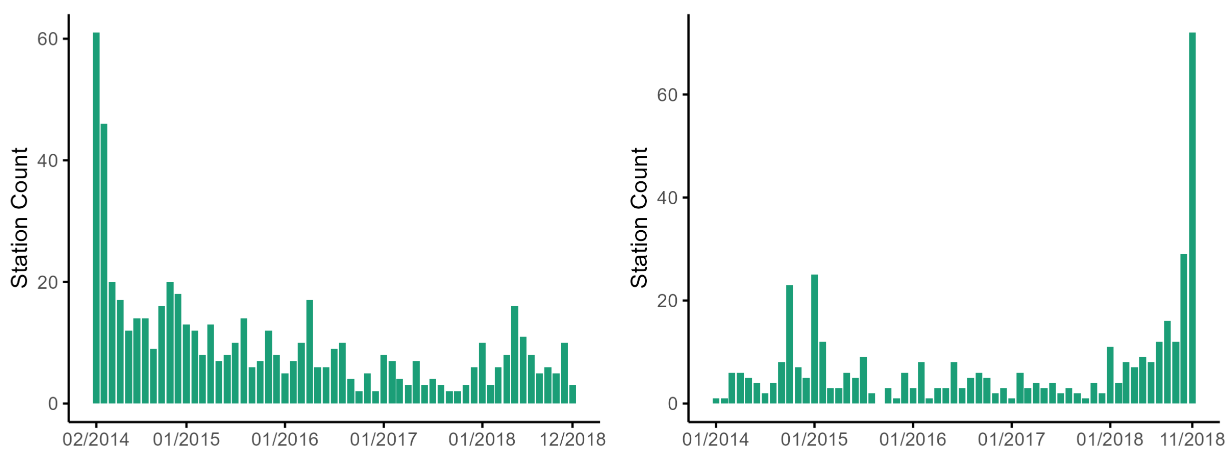
Notes: Data sourced from the Oil Price Information Service (OPIS)

Prices in OPIS are collected through monitored fleet credit card transactions, customer reports, and direct feeds from stations. OPIS reports prices for three gasoline grades—regular, mid-grade, and premium—which correspond to 87, 89, and 91 octane levels in California. OPIS reports a single price per station per day for each fuel type. However, a price is not observed every day for each station or each fuel blend sold. Panel (a) of Figure 1 shows the distribution of observation counts by station for the sample period 2014-2018 (1,825 days) for regular gasoline. Data coverage is high, with the median store having 1,756 prices reported. In fact, 75% of stores report more than 1,500 price observations over the sample period. To account for differential start and stop dates across stations during the sample period, Panel (b) of Figure 1 displays the number of price observations as a percent of potential reporting days for each station using the number of days spanned between the first and last price observations as the denominator. The median station reports prices for 97% of its respective sample period, and 75% of stations have prices reported on more than 86% of potential days.

2.1 Classification of Entering and Exiting Stations

To our knowledge, there does not exist an official public source of gas station operation dates in California.[7] We therefore leverage the high frequency of the OPIS data to identify entering or exiting stations. Conceptually, we classify entering or exiting stations based on the date of the first or last observed price.[8] We use the location IDs assigned by OPIS to track each station over time. The location IDs allow us to create a continuous series for each station, even if the station rebrands or is renovated.

Figure 2: Month of First and Last Observation by Station



(a) First Observation (b) Last Observation

Notes: Data sourced from Oil Price Information Service (OPIS)

For each station, we determine the date of the earliest price observation and the date of the last observed price during our sample period. Figure 2 graphs the number of stations by month of the first price observation in panel (a) and the number of stations by month of the last observation in panel (b). Panel (a) excludes January 2014 and panel (b) excludes December 2018 since the overwhelming majority of first and last observations fall in these two months, as to be expected.

[7]while business permit data or underground tank information can be used to locate stations, these sources lack the pertinent temporal component of on-site business operations which is the relevant metric of entry and exit for competing firms. State tax data on gas stations are reported at the owner level, and thus do not contain store-level information across multiple stores under the same ownership. State surveys conducted by the California Energy Commissions are annual, lacking any information on the date of entry or exit.

[8]In this design, stations that go dormant for a period but later resume reporting prices during the sample are not considered as an exit followed by a subsequent re-entry.

Classifying entry and exit based on the first and last date of reported prices creates three potential sources of measurement error. First, for entering and exiting stations, we might misestimate the exact date of entry or exit if the date of entry or exit does not align with the start or end of price reporting. Due to the high frequency and coverage of the OPIS data and the multiple ways in which the OPIS data is collected (through transaction data, crowd-sourced reporting, and brand partnerships), we think this is unlikely to significantly bias our results. After our calculated entry date or prior to our calculated exit date, the median station classified as entering or exiting reports prices on 78% of days.

Second, because prices are not reported for every station, every single day, a non-entering or -exiting station might appear to enter the sample after January 1, 2014 or leave the sample before December 31, 2018. Inappropriately classifying these stations as entrants and exits would attenuate our results. To mitigate the likelihood of false-positives, we classify a station as an entrant if the station has an initial price observation after March 1, 2014. Likewise, we classify a station as exiting the market if we last observe a reported price for the station before November 1, 2018.[9] This results in 484 unique stations having an entry date during the sample and 348 unique station exits.

Finally, there is also the possibility that a station goes unreported to OPIS and, consequently, never appears in our data. To assess this possibility, we validate the OPIS data by bench-marking the number of stations and the change in the number of stations over time against two other measures for California: The Census Bureau's County Business Patterns data and the estimated number of stations calculated by the California Energy Commission. The County Business Patterns data reports the number of businesses as of the week of March 12th of the appropriate year broken out by NAICS code. The data show a *net* increase of 228 gas stations in California from March 2014-March 2019. The CEC also undertakes an effort to estimate the number of stations in California based on returns from the A15 survey and other government data sources. For the same period, the CEC estimates a net increase of 190 stations in the state and around 10,000 gas stations in total,

[9] As a robustness check, we replicate our analysis using even more conservative bounds – six months from the start and end of our sample period and find qualitatively identical results.

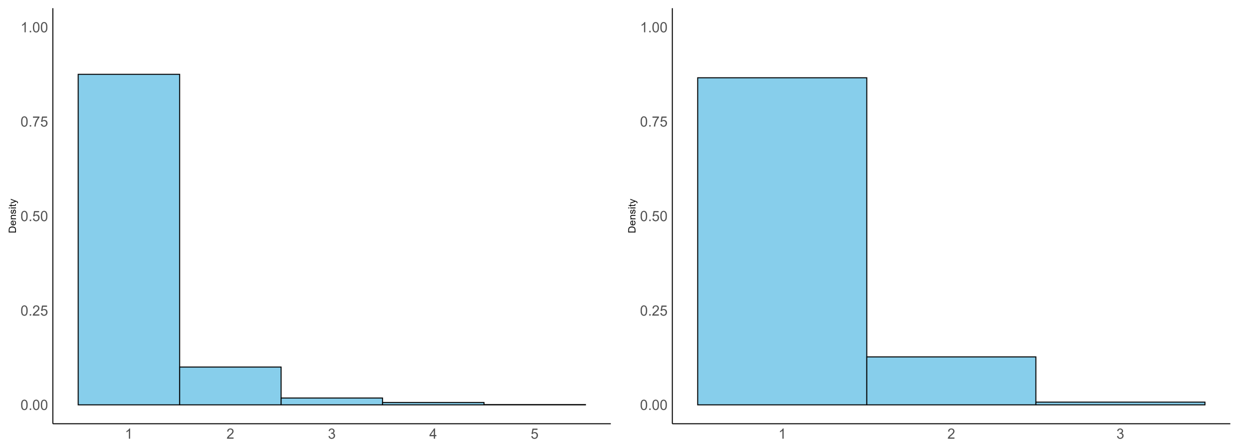
virtually identical to our calculations based on the OPIS data.

2.2 Patterns of Entry and Exit

Using the entry date, exit date, and geographic location for each station, we define the relevant market as the 1-mile radius circle around a station and calculate the number of competing stations at the station-date level.[10] Panel (a) of Figure 3 shows the number of entry events within 1 mile experienced by incumbent stations, conditional on experiencing at least 1 entry event. While 85% of stations do not experience a market entry during the sample period, experiencing multiple entries is also rare as 88% of stations that do experience entry only have 1 entrant during the sample period. Similarly, in Panel (b), 85% of stations do not experience a nearby exit, and 87% of those that do, only experience a single exit. To avoid contamination of our estimated effects from previous events, we focus on the first entry or exit event observed during the sample period in the subsequent analysis. Appendix Figure A2 shows the distribution of initial market sizes for the full sample of stations. Markets contain relatively few stations, with the median store having four other gas stations within a 1-mile radius. There are stations which operate in monopoly markets and one station located in downtown Los Angeles with 19 competitors within 1 mile. 10% of stations are located in markets with 8+ other stations. Appendix Figure A3 reports the number of entry events experienced by incumbent stations by their initial market size. The majority of entries occur in markets with 3-8 competitors and are the only entrant during the sample period. Finally, Figure A4 graphs the CDF for the distance between the incumbent station and the location of the first station entering or exiting within one mile - the distribution of distance is nearly uniform.

[10] A visual representation of the 1 mi. market definitions for Sonoma County, CA is shown in Appendix Figure A1. We also report results using other distance measures, varying from .5 to 10 miles.

Figure 3: Entries and Exits Experienced by Station: 2014-2018

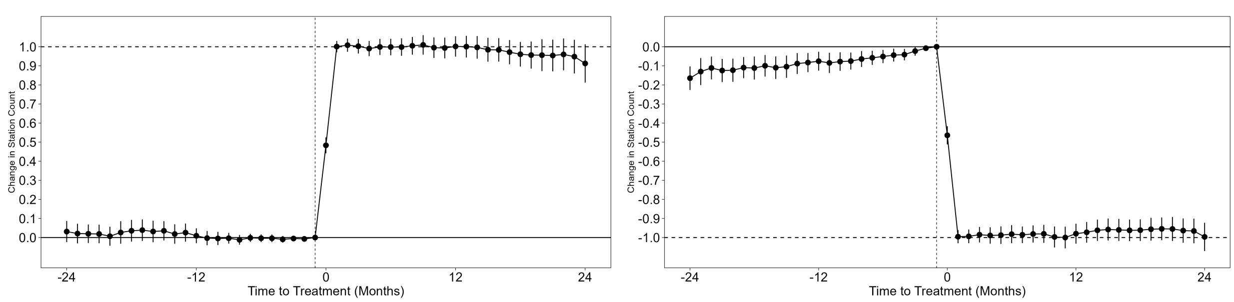


(a) Number of Entrants (b) Number of Exits

Notes: Data sourced from Oil Price Information Service (OPIS)

The causal estimates we present using both a difference-in-differences and event study design would be attenuated if the arrival of a gas station were simply a replacement for a nearby exiting firm within the same market. The same is true for an exiting station and subsequent entry. In the extreme, perfectly timed entry and exit would not contribute to identification as our measure of competition would not be impacted since the station count would remain constant across time.

Figure 4: Effect of Station Entry and Exit on Market Size



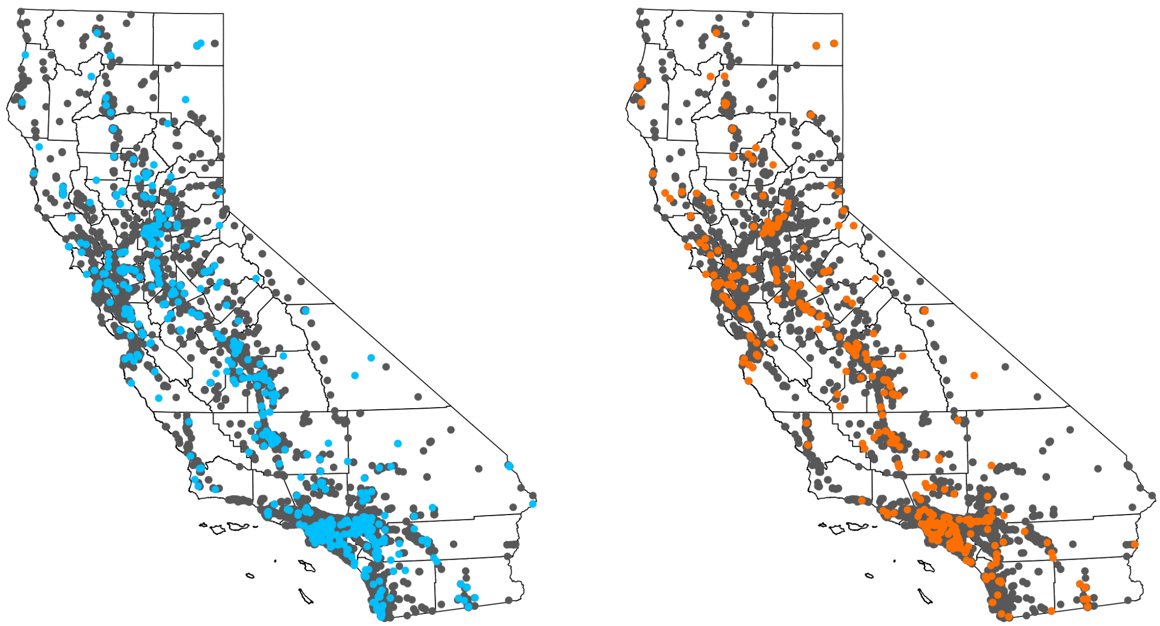
(a) Entry (b) Exit

Notes: Coefficient estimates for the effect of station entry (Panel (a)) and exit (Panel (b)) within 1 mile of an incumbent station on the number of competitors within 1 mile are shown. Event time t=-1 is the month prior to the event. Coefficients are estimated from a linear regression of station count on a panel of event time dummy variables, day-of-sample fixed effects, and station fixed effects. Standard errors are clustered at the city level. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

We can directly test for other correlated changes in the number of nearby stations by running an event study model, regressing the number of competitor stations within 1 mile on event time indicators for the first entry event and the first exit event, separately. We report estimated coefficients and standard errors in Figure 4 with entries in panel (a) and exits in panel (b). The regression includes station and day-of-sample fixed effects. Results suggest no evidence of exits preceding entry events, as indicated by precisely estimated zeros for the event-time coefficients leading up to the timing of entry. If there were significant exit preceding entry events, we would expect a downward sloping line above zero leading up to event-time zero as the market size decreases. At the time of entry, the station count increases by 1 eliminating simultaneous entry and exit as a source of bias. The change in market size is persistent over the post-period with minimal decay in the store count up to 1 year after the event, ruling out exit by the new station or another station, or additional entry in the post-period. This also aligns with the earlier descriptive analysis, which showed very few stations experienced multiple entries. Results for exit events are qualitatively similar, with only slight evidence of leading entries in the market. At the time of the exit, there is no evidence of simultaneous entry, and no evidence of lagging entry or subsequent exits in the year following the initial station exit event. With this evidence, we are confident that the identified entry and exit events represent a true shift in the competitive landscape.

Turning to geography, station entry and exit occurs throughout California. Figure 5 shows the location of the entrant gas stations in blue and exiting gas stations in orange. Although entry and exit are more concentrated in urban areas and along the I-5 and CA-99 corridors in the Central Valley, entry appears to largely match the locations of existing stations and not simply concentrated in new markets. We observe entry in remote and rural parts of the state, however these events will only contribute to identification of the parameters in the empirical analysis if they are located sufficiently close to an incumbent station.

Figure 5: Map of Station Entry and Exit: 2014-2018



(a) Entries

(b) Exits

Notes: Entrant stations are shown in Panel (a) by blue dots, with exiting stations represented by orange dots in Panel (b). Grey dots reflect incumbent stations that neither enter nor exit during the sample. Data sourced from the Oil Price Information Service (OPIS).

The main concern with previous cross-sectional analyses is that entry and exit are likely to occur in locations that differ from markets that do not observe a structural market change. Failure to account for these differences can lead to omitted variable bias. We test for baseline differences in observable characteristics, by combining demographic data from the 2014 American Community Survey's 5-year estimates with station characteristics taken from the OPIS data and the California Energy Commission's A15 survey.

In Table 1, we report the mean and standard deviation for a collection of variables at the station-level in the top panel, and at the tract level in the bottom panel. In both cases, we compare stations or tracts which enter (or have an entrant in the case of tracts) to exits with the difference in means and the accompanying p-value reported in Column 3. Focusing first on the station char-

Table 1: Baseline Differences By Entry or Exit Status

	Entry	Exit	Difference
Station Characteristics			
Station Count	484	348	
Branded Gas	0.250	0.345	-0.095
	(0.433)	(0.476)	(0.003)
Hypermart	0.052	0.023	0.029
	(0.222)	(0.150)	(0.026)
# of Competitors	3.479	4.569	-1.090
	(3.094)	(3.259)	(0.000)
Distance to Highway	1,323	796	527
	(3,041)	(1,423)	(0.001)
$<$ .25 Mile to Highway	0.510	0.563	-0.053
	(0.500)	(0.497)	(0.131)
Service Bay	0.013	0.065	-0.052
	(0.114)	(0.247)	(0.006)
Car Wash	0.160	0.119	0.041
	(0.367)	(0.325)	(0.191)
Convenience Store	0.660	0.662	-0.002
	(0.474)	(0.474)	(0.971)
Kiosk	0.052	0.060	-0.007
	(0.223)	(0.238)	(0.725)
Restaurant	0.082	0.060	0.022
	(0.274)	(0.238)	(0.338)
Supermarket	0.036	0.035	0.001
	(0.186)	(0.184)	(0.947)
Tract Characteristics			
Tract Count	435	319	
Gas Price	3.958	3.975	-0.016
	(0.152)	(0.178)	(0.210)
Income (Median)	55,182	55,626	-444
	(24,118)	(25,880)	(0.811)
Income (Mean)	69,056	72,416	-3,360
	(27,739)	(35,018)	(0.158)
Poverty Rate	19.912	19.996	-0.084
	(12.403)	(13.100)	(0.929)
Households	1,957	1,921	36
	(958)	(922)	(0.600)
House Value (Median)	275,643	330,559	-54,916
	(177,797)	(228,081)	(0.000)
% No Vehicle	7.024	8.381	-1.358
	(6.513)	(8.210)	(0.015)
% Commuting by Vehicle	86.245	84.977	1.268
	(10.509)	(11.208)	(0.116)
Commute Time	26.812	25.495	1.317
	(7.112)	(6.113)	(0.007)

Notes: Means and standard deviations for station-level and tract-level characteristics are reported in Columns 1 and 2. The difference in means is shown in Column 3 with the associated p-value reported below. Data are sourced from the 2014 American Community Survey 5-year estimates, the Oil Price Information Service, and the California Energy Commission.

acteristics, we report that entering stations are more likely to sell unbranded gasoline and to be a hypermart than exiting stations. They also have fewer competitors within the 1 mile market. In terms of station amenities, entrants have fewer attached service bays for repairs, but more car washes and restaurants than exiting stations.

Turning toward tract characteristics, we see that gas prices in the first 3-months of our sample in tracts that experience an entry are indistinguishable from prices in tracts that experience exit. Census tracts that experience entry have similar household income and poverty rates, and housing density as tracts that experience station exit. However, tracts where exits occur have higher house values. Focusing on related driving characteristics, we see that tracts with entry have lower rates of households with no vehicle and higher rates of people commuting via vehicle with longer commutes.

The baseline demographic differences between locations that do and do not observe market size changes highlight the need to account for the inherent market characteristics to address the endogenous entry, exit, and continuing operation decisions of stations. Additionally, to the extent that there are unobservable demographic characteristics that are correlated with both station entry and demand for gasoline, cross-sectional regressions of price on the number of competitors are likely to yield biased estimates of the effect.

3 Empirical Strategy

The localized nature of gasoline station competition allows for the geographic and temporal variation in exposure to station entry and exit across firms to form the basis of a difference-in-differences estimation for the causal effect on incumbent pricing. Following prior work by Arcidiacono et al. (2020), we treat the exact timing of the entry or exit of a new gasoline station as a short-run exogenous shift in the market structure for incumbent firms after conditioning on the inherent market structure.

Importantly, we include a rich panel of fixed effects to account for unobserved variable bias inherent to the endogenous location decision of entering and exiting firms. By restricting the model to identification from within station variation in the number of nearby competitors, over time, the model accounts for factors important to the location decision such as the overall price level in the market, local price elasticity of demand, local traffic patterns, and relevant customer characteristics.

Formally, we estimate:

$$P_{st} = \alpha + \beta N_{st} + \sigma_s + \delta_t + \Phi_c(t) + \varepsilon_{st} \tag{1}$$

where the main outcome variable is the retail price in dollars per gallon at station s on day t and N_{st} is the count of competitors to station s on day t, increasing upon entry and decreasing with nearby exit. In our preferred specifications, we define the relevant market as the 1-mile radius circle around the incumbent station, consistent with prior literature (Lewis 2015; Davis et al. 2023; Fischer et al. 2023; Hastings 2004; Bernardo 2018; Carranza et al. 2015; Barron et al. 2004). We focus on the price of regular-grade unleaded gasoline, which in 2018 accounted for roughly 70% of retail sales in California.

Station fixed effects (σ_s) are included to capture time-invariant differences between locations, such as station amenities and size, location effects, and distance to the wholesale terminal which largely drives differences in input costs. Day-of-sample fixed effects (δ_t) capture state-wide daily shocks to both input costs, such as oil prices and refinery supply shocks, as well as common daily shocks to product demand. Lastly, in the preferred specifications, we include a city-specific linear time trend, $\Phi_c(t)$, to account flexibly for city-level trends that may be correlated with price and station demand. We cluster standard errors at the city-level to account for common shocks across units.

The specification above constrains the impact of entry and exit on prices to be the same magnitude magnitude (although of opposite sign). Relaxing this assumption, we also estimate the following static difference-in-differences specification:

$$P_{st} = \alpha + \mathbb{1}Entry_{st} + \sigma_s + \delta_t + \Phi_c(t) + \varepsilon_{st}, \tag{2}$$

and the corresponding specification for station exits. The specification replaces the station count variable with an indicator variables for the first entry or exit experienced by the incumbent.

Equations (1) and (2) are two-way fixed effects (TWFE) estimators, in which stations that do not experience a change to market size within 1 mile during the sample period and previously treated stations both serve as control units for stations that experience entry or exit. To address the potential bias of TWFE estimators in a setting with staggered, heterogeneous treatment, we implement the regression-based estimator from Gardner et al. (2024) that provides results that are robust to heterogeneous, staggered treatments while providing similar confidence intervals to standard TWFE estimators if treatment effects are homogeneous. In contrast, alternative estimators rely on estimating effects for each cohort immediately before and after treatment non-parametrically to ensure appropriate control units are used in estimation. These approaches are computationally inefficient in the presence of many treated cohorts, defined by the exact date of entry in our current data specification. Additionally, to the extent that there is any mismeasurement of the exact date of entry or exit, these methods can yield biased estimates.

The approach in Gardner et al. (2024) regresses the outcome, gas prices in our setting, on group and period indicators using only untreated and not-yet-treated observations. In the second stage, the previously estimated group and period effects are subtracted from the outcome variables to create a new residualized outcome variable which is then regressed on the event-time treatment variables. This results in familiar event-study coefficients comparable to the results from the two-way fixed effects approach.

Identification of the main coefficient of interest, β, as the causal effect of a change in the number of nearby stations on prices requires two main assumptions. First, the main identifying assumption requires entry and exit to be conditionally uncorrelated with the error term. Specifically, conditional on station fixed effects, time fixed effects, and time trends, the changes in station count are exogenous.

$$E[\varepsilon_s|\sigma_s, \delta_t, \Phi_c(t), N_{st}] = 0 \tag{3}$$

This requires that there were no other factors correlated with the timing of the station entry or exit that also impacted the pricing of nearby stations. Although this assumption cannot be directly tested, demonstrating that the treatment and control follow common trends in the pre-treatment period offers a falsification test of the assumption. Secondly, the differences-in-differences framework assumes stable unit treatment values which requires that there is no spillover of treatment onto control units outside of the impacted market. This is plausibly satisfied in our setting due to the local geographic nature of gas station competition, limiting the spillover price effects from treatment units to the larger pool of control observations.

To better examine the price dynamics of entry and exit, and to document the lack of differential pre-treatment trends, we estimate the following event study model for station entries, and the equivalent analog for exit events separately:

$$P_{st} = \alpha + \sum_{k=-24}^{24} \beta_k \mathbb{1}[Entry_{st} = k] + \sigma_s + \delta_t + \Phi_c(t) + \varepsilon_{st} \tag{4}$$

setting event time indicators for the number of months before and after the first nearby entry or exit observed at station s. End points are binned to include 24 or more months before/after the event.

The event study approach offers a complement to the difference-in-difference specifications in equations (1) and (2). Notably, the event-study specification offers direct evidence on whether pre-treatment trends are parallel for treated and control stations. If the prices in markets that observe entry or exit follow a different trend over time than the "control" locations, differential trends would bias the difference-in-difference estimators. Precisely estimated null effects in the time periods leading up to the entry or exit event provide support that treatment and control markets were following common trends. In addition, the event study specification allows for the effect of entry or exit to evolve dynamically, illustrating whether prices change quickly or gradually following a change in the competitive landscape. As before, we estimate event studies using the estimator from Gardner et al. (2024).

The models specified above provide estimates of the average treatment on the treated (ATT) when the identification assumptions are satisfied. Given the baseline differences in treatment and control areas shown in Table 1, the estimated coefficient represents an internally valid estimate of the causal effect of entry or exit at locations where firms decide to enter or exit.

4 Results

4.1 Station Count Results

We present the estimates for the relationship between nearby station counts and the pricing of incumbent stations in the top panel of Table 2. Column 1 presents the coefficients estimated from the simple linear regression of price (in dollars per gallon) on the station count variable as an initial point of comparison. Columns 2 - 4 add station fixed effects and day-of-sample fixed effects. Our preferred specification in column 5 further adds city-specific linear time trends.

In column 1, we estimate an economically small, but statistically significant negative relationship between market size and prices. The addition of an additional station within one mile lowers prices at incumbent firms by 0.6 cents. However, this specification fails to account for the endogenous relationship between price and entry/exit, leading the estimate to likely be biased.

In column 2, we include station-level fixed effects to control for market characteristics plausibly correlated with station density. In this specification, identification arises from within-station variation. An additional nearby competitor is now associated with a significant 6.7 cent decrease in prices. Relative to column 2, the attenuation of the estimates in column 1 is consistent with the hypothesis that station density is endogenously higher in high-price regions, leading to a likely bias in the estimate.

In column 3, we now add day-of-sample fixed effects to the naive regression to account for daily location-invariant shocks that affect all stations, such as changes in input costs, refinery outages and major weather events. Unsurprisingly, day-of-sample fixed effects account for roughly 80 percent of the variation in prices.

Column 4 is the canonical two-way fixed effects model including both day-of-sample fixed effects and station fixed effects. With the inclusion of both station and day-of-sample fixed effects,

Table 2: Effect of Changes in Competition on Incumbent Pricing

	(1)	(2)	(3)	(4)	(5)	(6)
Station Count	-0.006***	-0.067***	-0.005**	-0.012***	-0.010***	
	(0.002)	(0.021)	(0.002)	(0.003)	(0.003)	
R Sq.	0.001	0.163	0.809	0.956	0.958	
Obs.	14,749,777	14,749,777	14,749,777	14,749,777	14,749,777	
Entry	-0.128***	-0.310***	-0.044***	-0.023***	-0.020***	-0.025***
	(0.011)	(0.023)	(0.010)	(0.004)	(0.004)	(0.004)
R Sq.	0.005	0.170	0.809	0.956	0.958	
Obs.	14,749,777	14,749,777	14,749,777	14,749,777	14,749,777	14,748,218
Exit	-0.066***	-0.193***	-0.003	0.002	0.001	0.000
	(0.014)	(0.030)	(0.011)	(0.004)	(0.004)	(0.005)
R Sq.	0.001	0.165	0.808	0.956	0.958	
Obs.	14,749,777	14,749,777	14,749,777	14,749,777	14,749,777	14,749,777
Station FE	No	Yes	No	Yes	Yes	Yes
Day of Sample FE	No	No	Yes	Yes	Yes	Yes
Linear Time Trend	No	No	No	No	Yes	No
Estimator	TWFE	TWFE	TWFE	TWFE	TWFE	DID2S

Notes: The table reports estimates for the effect of a change in nearby competitors within 1-mile on incumbent pricing for regular unleaded gasoline in dollars per gallon. The top panel uses the daily count of nearby stations as the independent variable. The panels for Entry and Exit use an indicator variable for the dates after the first nearby entry or exit. Column 1 reports estimates from a simple linear regression. Columns 2-5 report results with the addition of the fixed effects listed in the panel below. Column 6 uses the two-stage DID estimator from Gardner et al. (2024). Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018. Model standard errors are reported in parentheses and clustered at the city level. *** = significant at 1 percent level, ** = significant at 5 percent level, * = significant at 10 percent level.

we estimate an additional competitor reduces prices by roughly 1.2 cents per gallon. Relative to the larger point estimate in column 2, the inclusion of day-of-sample fixed effects in addition to station fixed effects controls for statewide trends in prices and station density. Since, on net, station density has increased and prices fell over our study period, the coefficients in column 2 are biased downwards relative to the two-way fixed effects estimates in column 4.

Column 5 represents the preferred specification, showing results are robust to the further inclusion of a city-specific linear time trend. An additional competitor within one mile results in a 1.0 cent reduction in incumbent prices, which represents around a 2.5% reduction in firm markups during the sample period which average 40 cents.[11] These estimates control for potential selection if entry or exit are correlated with local demand or price trends. The robustness of the coefficient to the inclusion of a city-specific linear time trend between columns 4 and 5 provides support that treatment and control markets are not trending differentially.

4.2 Effects of Entry and Exit

In the second and third panels of Table 2, we present results from estimating equation (2), regressing incumbent prices on indicator variables that reflect the period after the first entry or exit within one mile faced by the station. As with the results based on station counts, column 1 presents the results from a bivariate regression that omits fixed effects and columns 2 - 5 successively add station fixed effects, day-of-sample fixed effects, and city-specific linear time trends.

We focus attention on columns 4 and 5 that control for both time-invariant and station-invariant unobservables and identify the coefficient on the entry and exit indicators from within-station variation relative to statewide trends in prices. Here, we find that the earlier estimated impacts from the regression using changes in nearby station count operate entirely through the impacts of entry. In our preferred specification, the pricing of an incumbent firm falls by 2.0 cents per gallon following the entry of a new competitor within 1 mile. In contrast, we find little evidence that incumbent pricing changes following the exit of a competitor, estimating a precise null

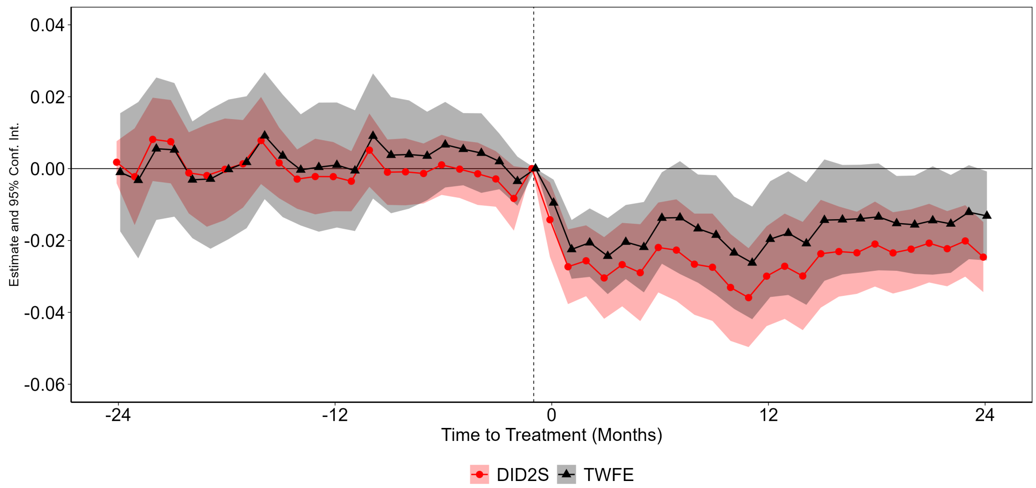
[11] California Energy Commission estimates of CA gasoline price breakdown and margins.

effect. By breaking out the entry and exit effect separately, we can see that the prior results using the nearby station count as the regressor, which uses variation from both entry and exit events for identification, attenuated the effect due to asymmetric effects by event type. In column 6 we show that results for both entry and exit are robust to estimation using the estimator from Gardner et al. (2024) to account for the staggered treatment timing.

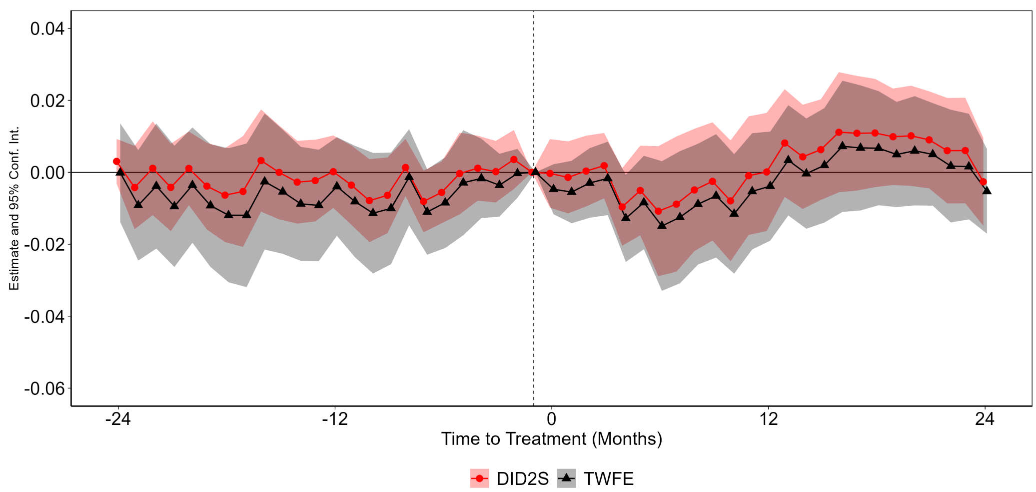
The event study specification, formalized in equation (4), allows us to examine the speed with which incumbent prices change after the entry or exit of a nearby competitor, and post-event pricing dynamics over time. In addition, the event study design allows us to visually and statistically assess the assumption that the treatment and control evolved along common trends in the pre-treatment period. This provides support for the identifying assumption of the difference-in-differences framework, that prices for control and treated stations would have evolved along similar patterns, in the absence of the treatment.

Figure 6a presents results for the first entry event experienced by a given incumbent station, and Figure 6b reports results for the first exit event experienced during the sample period. We plot the coefficients from the Gardner et al. (2024) estimator in red, as well as the coefficients of the canonical TWFE estimator in black.

Figure 6: Effect of Station Entry and Exit Within 1 Mi. on Incumbent Pricing-DID2S



(a) Station Entry



(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in Panel A) and station exit (in Panel B) within 1 mile of an incumbent station on incumbent pricing of regular unleaded gasoline are shown. Results from the canonical TWFE estimator are shown in black. Results from the Gardner et al. (2024) two-stage DID estimator are shown in red. Data are sourced from the Oil Price Information Service (OPIS).

Examining the pre-treatment point estimates, we see that prices for treated and control stations were parallel prior to treatment. Point estimates for the event-time coefficients for the months prior to both entry and exit events are close to zero and all time periods include zero within the confidence intervals. The null estimates consistent across periods before entry also provide supporting evidence that incumbents do not lower prices in anticipation of entry. Rather, incumbents only drop prices at the start of new operations and accommodate entry. Likewise, the lack of an effect in the periods before exit provides evidence that stations do not successfully engage in predatory pricing to force the exit of a nearby station.

Examining the post-treatment coefficients for entry (in Figure 6a), we see that the price set by an incumbent stations falls discretely and immediately after the entry of a new, nearby competitor. The entry of a station is associated with a sharp drop in the price at incumbents of 2.7 cents. Effects are precisely estimated and persistent over the long run suggesting that the entry of a new station results in a quick shift to a new, lower price equilibrium. Figure 6b shows precisely estimated null effects for exit events which stand in contrast to the negative estimated effects of entry. In both cases, we see that the results from the robust estimator from Gardner et al. (2024) perform better than the TWFE estimator; we see tighter confidence intervals around 0 in the pre-treament periods, and stronger evidence of a persistent negative price effect upon entry. As such, in the event studies that follow, we only report results using the estimator from Gardner et al. (2024).

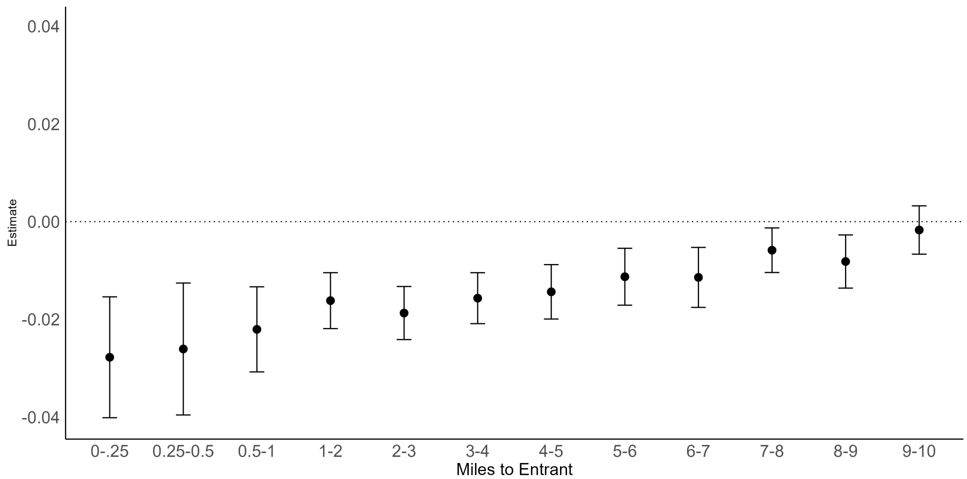
4.3 Do the effects of entry and exit vary by proximity and fuel grade?

We next explore whether the effects of entry or exit on the pricing of incumbent firms vary on two dimensions: (1) geographic proximity to the entrant, and (2) by grade of gasoline. Gasoline stations compete in a geographically differentiated market with nearby stations in closer competition than more distant stations (Houde (2012), Chandra and Tappata (2011), Eckert and West (2005)). If entrants impose a competitive impact on the pricing of incumbent firms, we would expect, all else equal, for the effects to be greatest when an incumbent faces nearby entry rather than a more distant entrant. Similarly, the exit of a more distant competitor should have reduced effects. In addition, stations sell different grades of gasoline. As noted above, regular unleaded gasoline (with an octane level of 87), represents around 70% of motor gasoline sales in California. Premium (91 octane) accounts for roughly 25% of motor gasoline sales, with mid-grade accounting for the remainder. Building on past work that finds variation in the elasticity of demand, by grade of gasoline, (Yatchew and No (2001)), as well as evidence of substitution between grades of gasoline (Hastings and Shapiro (2013)), entry or exit might differentially affect incumbent pricing for each grade of gasoline.

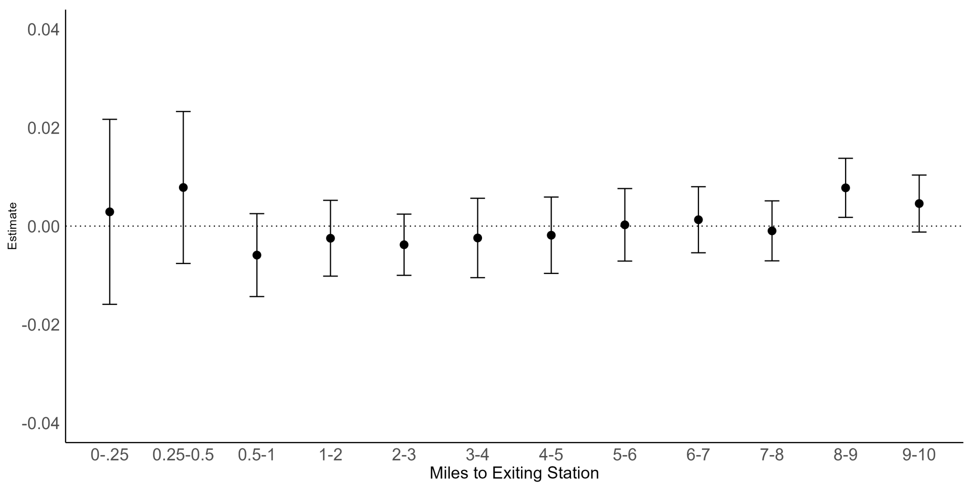
To estimate variation in the effects of entry by proximity, we extend the specification in equation (2). As with the earlier specification, we proxy for station competition using the straight-line distance between an incumbent station and an entrant. Although this abstracts from the nuances of local road networks and commuting flows, leveraged in Houde (2012); Davis et al. (2023), this approach can be easily implemented for all incumbent stations state-wide.

We run regressions for each distance bucket separately designating treatment as the month of the first entry/exit that occurs between the minimum and maximum distance for the bucket. We present the point estimates and standard errors from the regression using the two-stage estimator in Gardner et al. (2024) for entry in Figure 7a and exit in Figure 7b. As with the estimates in Table 2 derived from equation (2), stations that do not experience any entry or exit serve as controls.

Figure 7: Effect of Station Entry and Exit on Incumbent Pricing, by Distance



(a) Station Entry



(b) Station Exit

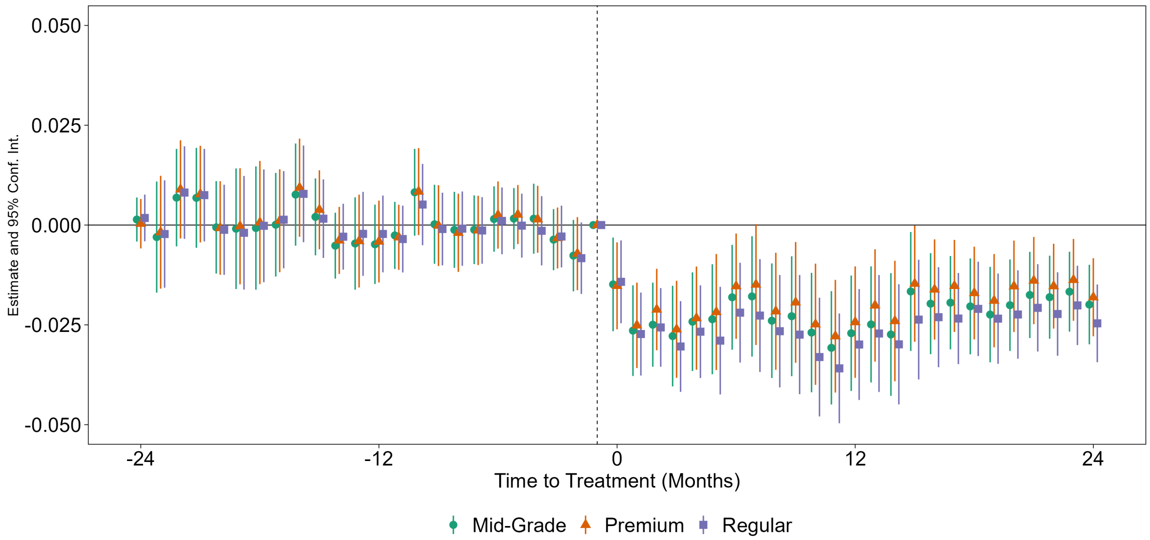
Notes: Coefficient estimates for the effect of station entry (in Panel A) and exit (in Panel B) by distance from incumbent station on incumbent pricing of regular unleaded gasoline are shown. Coefficients are estimated by running separate regressions by distance bucket using the two-stage DID estimator in Gardner et al. (2024). Data are sourced from the Oil Price Information Service (OPIS).

Results vary by distance, as expected, with the strongest entry effects observed for events occurring within a quarter-mile of the incumbent. The effect monotonically increases towards zero as the distance from the station increases, with economically irrelevant effects after 7 miles. This is consistent with prior research which shows retail gas competition is highly localized. We again estimate null effects for the exit of a nearby station across all distances.

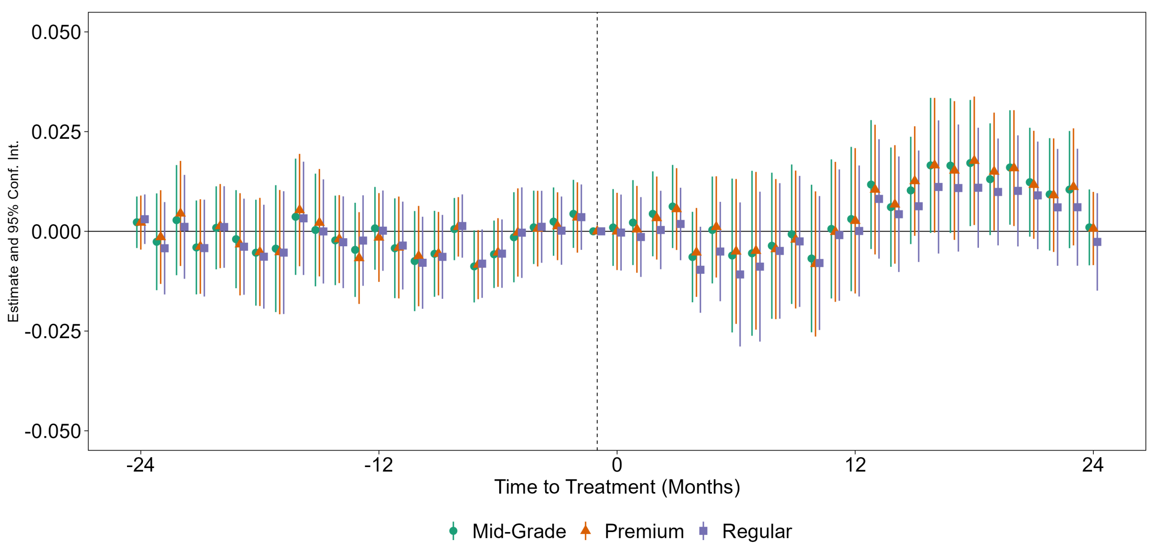
Identification of causal estimates in our setting relies on the SUTVA assumption. Our distance results show that this assumption is valid for most distances with limited spatial spillovers of entry. As a robustness check, we can exclude control stations which do not have entry within 1 mi., but do experience an entry within the 7 mi. cutoff for significant effects shown above. Appendix Figure A5 reports the coefficients and shows that estimated results are even stronger at the time of entry, persist at the increased level, and still satisfy the pre-trends assumption. Confidence intervals remain similar in magnitude despite the reduction in sample size.

Second, we estimate the effects of entry and exit separately for the different grades of gasoline. We present estimates for the effect on incumbent pricing for all three grades for entry in Figure 8a and exit in Figure 8b. We find qualitatively consistent results across the three grades. For entries, point estimates for the price decrease are steepest for regular gasoline, followed by mid-grade and then premium at each post-treatment time period. However, we cannot statistically reject that all three blends have the same coefficient. As was the case for regular gasoline, exit events are not associated with statistically distinguishable changes to incumbent pricing for either of the other two blends.

Figure 8: Effect of Station Entry and Exit on Incumbent Pricing, by Blend



(a) Station Entry



(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in Panel A) and station exit (in Panel B) within 1 mile of an incumbent station on incumbent pricing by gasoline blend type. The regression is estimated using the two-stage DID estimator in Gardner et al. (2024). Data are sourced from the Oil Price Information Service (OPIS).

5 Asymmetry of Entry and Exit Effects.

In the preceding section, we presented evidence that prices at incumbent stations decline significantly and immediately following the entry of a competing station. The impact on incumbent pricing attenuates with the distance to the new competitor and the effects are largely consistent across different grades of gasoline. In contrast, we find little evidence that station exit causes prices to rise, finding consistent, precisely estimated null effects across distances and grades. Although models of symmetric competition largely suggest symmetric impacts of entry and exit on incumbent pricing, we explore four potential sources of heterogeneity that help to explain the difference in the effect of entry and exit on incumbent pricing, and, in particular, the null effects observed for exit.

First, as we noted when discussing Table 1, stations that exit the market tend to exit from locations that are systematically different than those chosen by entering stations. A long literature (e.g., Seim (2006)) highlights the geographic element of endogenous entry decisions – firms choose to enter markets in which they can more easily differentiate themselves from their competitors. In our setting, we observe that entry is more likely to occur in locations with fewer competitors and further from highways. To the extent that heterogeneity exists in the magnitude of the effects of entry and exit that is correlated with location, the null results for exit might be partially explained by heterogeneity in the types of locations that stations exit compared to the locations that stations enter.

Second, entering and exiting stations may differ on observable and unobservable dimensions that impact competitiveness and the degree to which their entry or exit might impact incumbent pricing. Observably, entering stations differ from those that exit in ways that impact competitiveness. As noted earlier, stations that exit are more likely to be branded than stations that enter. If the competitive effect of a station is systematically correlated with branding, the null effect we observe for exiting stations might be explained by compositional differences in the types of stations that enter and exit markets.

Stations may also vary on unobservable characteristics correlated with the competitiveness or attractiveness of the stations. In a model of endogenous exit with heterogeneous costs or productivity (e.g., Asplund and Nocke (2006)), stations with idiosyncratically higher costs or lower productivity will tend to have lower profits and, all else equal, be more likely to exit the market. If, prior to exit, these stations tend to exert less competitive pressure on neighboring firms, their exit might not impact prices to the same degree as an entering station.

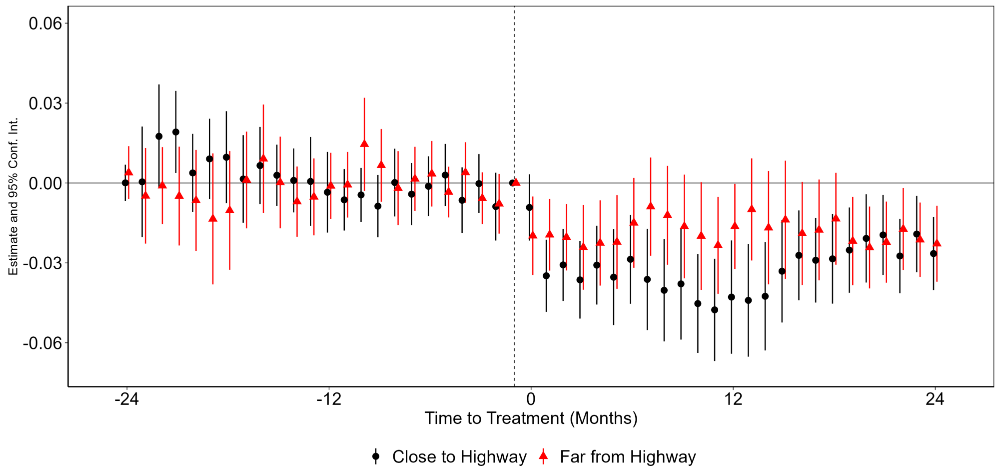
Finally, we consider mismeasurement of exiting firms as a source of attenuation bias that might explain the null result observed for exit. Although our counts of entering and exiting firms largely align with administrative counts from the Census Bureau and California Energy Commission, if stations with continuing operations lack (for whatever reason) reported prices during the last two months of our sample, we would mis-classify them as exiting stations.

5.1 Geographic heterogeneity in the location of entry and exit

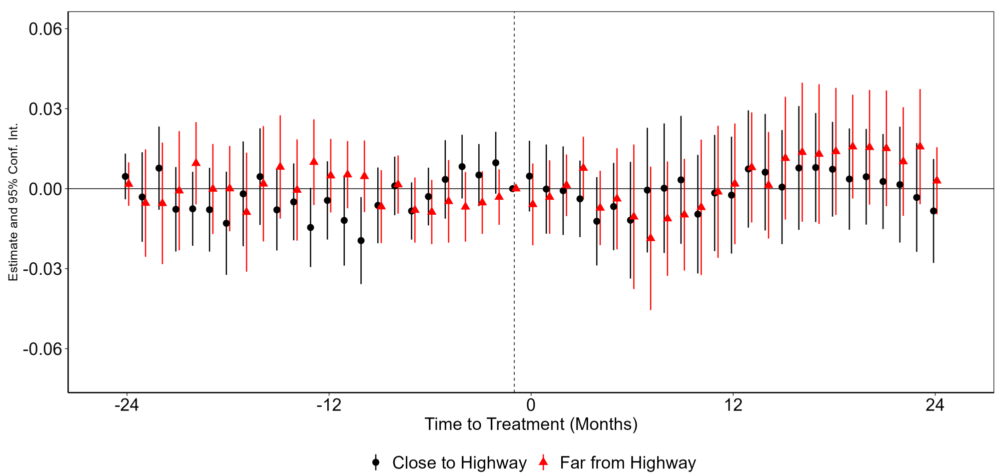
In Table 1, we noted that the locations where stations enter and exit are systematically different, varying with proximity to highways and the density of nearby stations. We consider the compositional differences in the markets in which stations enter and exit, first by estimating the event study model separately for entry and exit events near and far from highways. Specifically, we compare the effect of an entry or exit within a quarter mile of a highway on incumbents, to events that occur further away.

In Figure 9a, we observe modest differences in the effect of entry. Incumbent stations that face new competition from entrants located within 1/4 of a mile of a highway lower prices significantly more than do incumbent stations facing competition from an entrant located far from highways. Notably, although we find some heterogeneity when splitting the sample by the highway proximity for entering stations, the composition differences would tend to reduce, rather than augment the asymmetry we find – entering stations are more likely to be located *farther* rather than near to highways than exiting stations. As before, we fail to find a meaningful heterogeneity in the impact of exiting firms, splitting the sample by the highway proximity and plotting the estimated coefficients in Figure 9b.

Figure 9: Effect of Station Entry and Exit, by Distance to Highway



(a) Station Entry



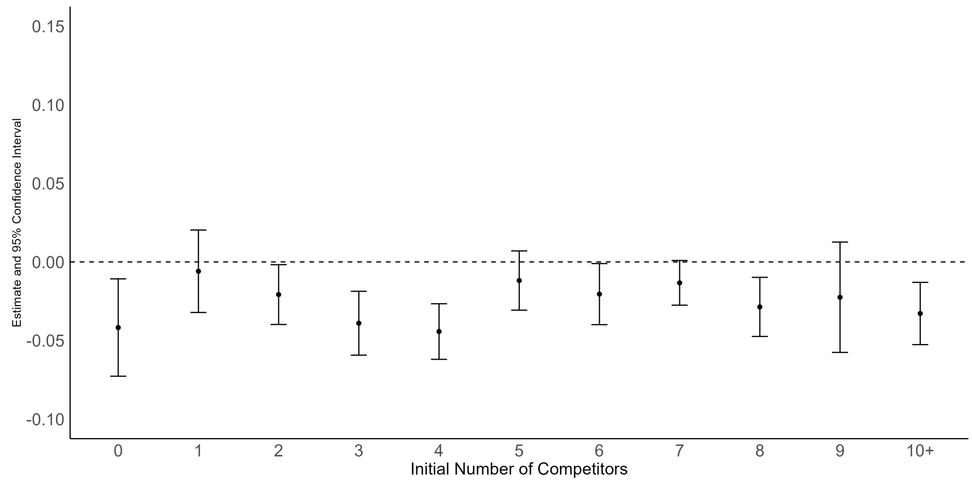
(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in Panel A) and station exit (in Panel B) within 1 mile by distance from the entering/exiting station to the nearest highway on incumbent pricing of regular unleaded gasoline are shown using the two-stage DID estimator in Gardner et al. (2024). Data are sourced from the Oil Price Information Service (OPIS).

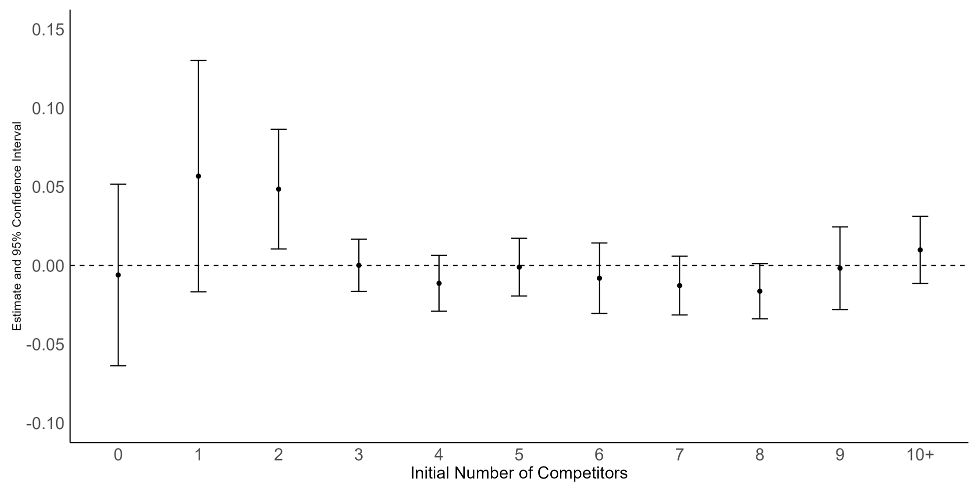
There is not a consensus in the literature on the effect of changes to market size on price levels in gas markets. Canonical work by Bresnahan and Reiss (1991) on competition in homogeneous goods markets suggests that entry into smaller, consolidated markets results in larger competitive effects and that this effect dissipates as the number of competitors in a market increases. However, Armstrong and Vickers (2022) and Barron et al. (2004) show that this result can reverse in search models with price dispersion depending on the search costs involved. We test for heterogeneous effects of entry and exit across the number of competitors in our setting by estimating equation 2 separately by the number of competitors faced by the incumbent station.

Figure 10 plots the coefficients for entry (in panel A) and exit (in panel B) based on the number of competitors faced by the incumbent station at the beginning of our sample. In Panel A, we find that incumbent stations facing zero, two, three or four competitors within one mile lower prices upon entry of an additional competitor. These cases make up the majority of the entry events we observe – roughly two-thirds of the incumbents affected by entry face fewer than five competitors at the start of the sample period. We see little evidence that entry impacts the prices of incumbent stations that initially face more than four competitors. This, perhaps by coincidence, aligns with the findings in Bresnahan and Reiss (1991) which found little competitive impact on incumbent firms of entry once five competitors were present in a local market. In retail gasoline markets, Tappata and Yan (2017) finds a similar threshold of market size for entry effects.

Figure 10: Effect of Station Entry and Exit on Incumbent Pricing, by Initial Market Size



(a) Station Entry



(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in Panel A) and exit (in Panel B) within 1 mile of an incumbent station on incumbent pricing of regular unleaded gasoline are shown using the two-stage DID estimator in Gardner et al. (2024). Coefficients are estimated by running separate regressions by the number of initial competitors faced by the incumbent. Data are sourced from the Oil Price Information Service (OPIS).

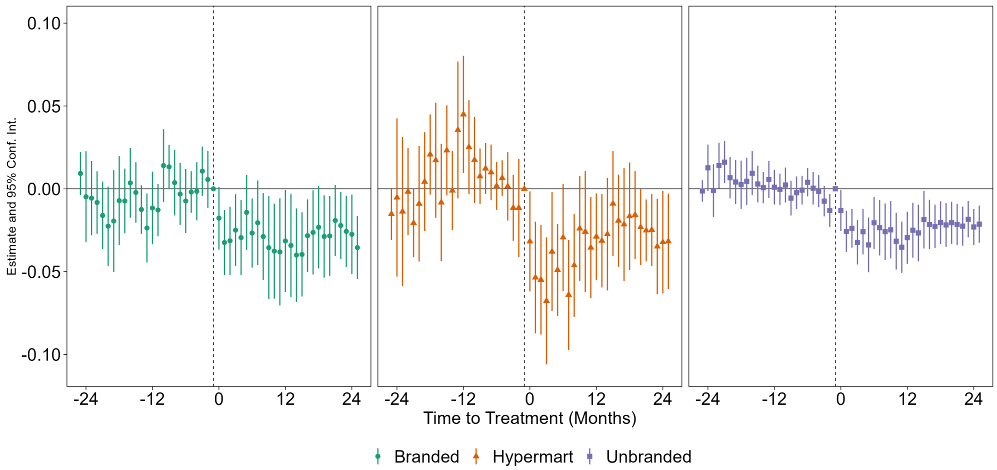
For exit events, we do find a relationship between market concentration and the impact of exit on incumbent pricing with some positive effects for incumbent firms facing relatively few initial competitors. Notably, point estimates suggest that incumbent stations with one or two competitors raise prices by roughly five cents per gallon when one of the competitors exits, although only the estimate for the latter is statistically significant. Yet, these cases account for a relatively small fraction (15%) of the roughly 350 exit events in the sample. We find robust null effects of exit for the remaining 85% of cases, the vast majority of which are settings in which the incumbent faced competition from more than three stations prior to the exit event. As a point of comparison, incumbents with one or two competitors comprised a larger fraction (roughly 25%) of those affected by entry.

5.2 Heterogeneity in entering and exiting station characteristics

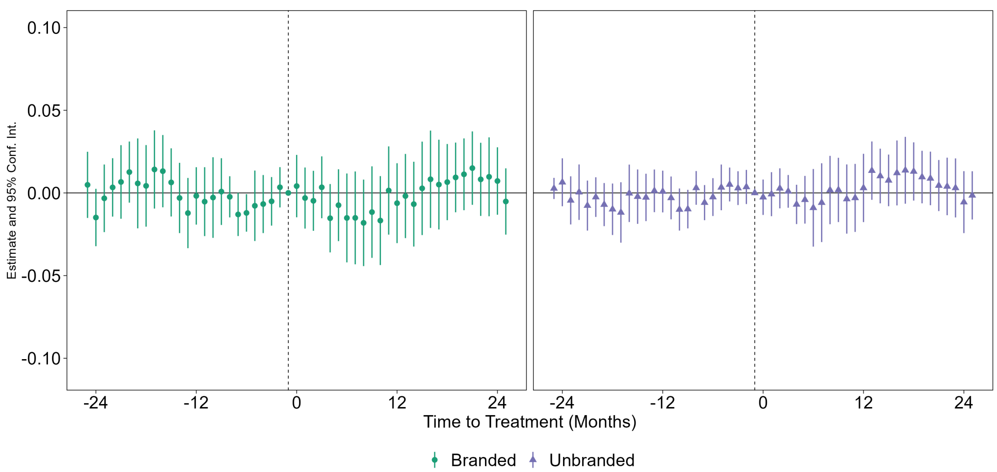
We next examine whether the asymmetry in the effects of entry and exit can be attributed to compositional differences in the types of branding of entering and exiting stations. Comparing entering and exiting stations in Table 1, entering stations are less likely to sell branded gasoline. Traditionally, unbranded gasoline is sold at a discount compared to branded gasoline. This heterogeneity in pricing can lead to differential effects of competition. Additionally, there has been an increase in the entry of high-volume stations referred to as hypermarts, such as Kroger, Costco, and Sam's Club. These stations sell unbranded gasoline and are characterized by having numerous pumps and high sales volume, further contributing to potential differences in competitive dynamics.

In Figure 11a, we test for heterogeneous entry effects for stations that sell unbranded vs. branded gasoline at the time of their entry and effects for hypermart entries only. Entry of a nearby unbranded station results in the familiar immediate 2-cent reduction in incumbent prices while the estimated effect for branded stations is slightly lower at 3 cents, however due to the standard errors for the estimates, we cannot reject statistically similar effects. For the entry of a nearby hypermart, the estimated entry effect is greater, at 5 cents on incumbent stations. In Figure 11b, we test for heterogeneous exit effects by the station's gas branding. Here, we omit the results for hypermarkets as only a handful of hypermarts exit the sample. As with our baseline results, we find precise null exit effects for both branded and unbranded stations.

Figure 11: Effect of Station Entry and Exit, by Station Category



(a) Station Entry

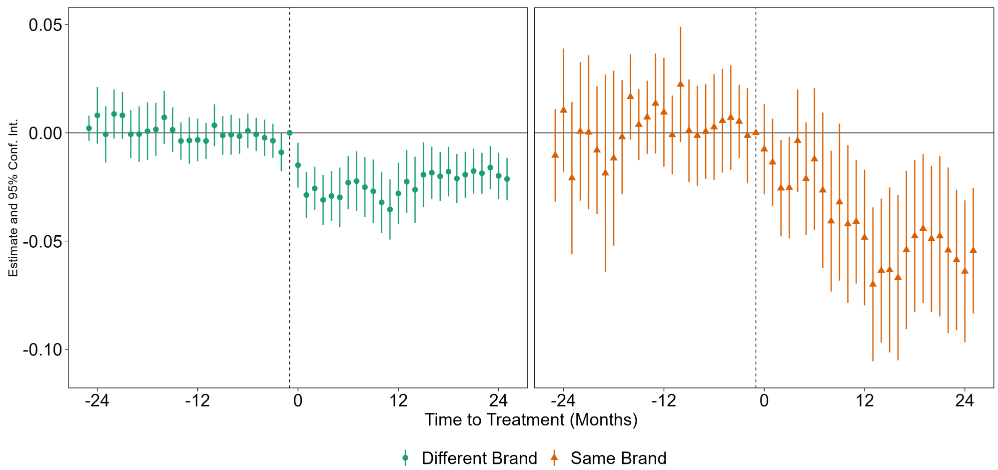


(b) Station Exit

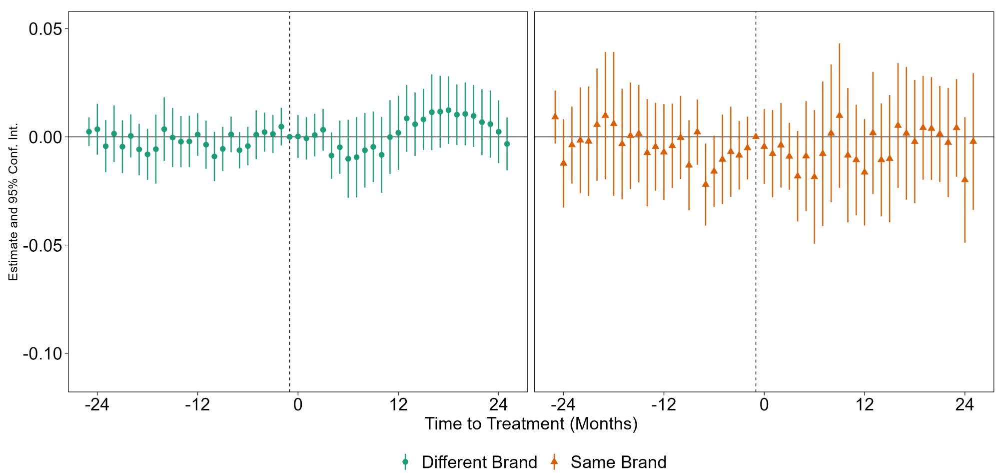
Notes: Coefficient estimates for the effect of station entry (in Panel A) and exit (in Panel B) within 1 mile on incumbent pricing of regular unleaded gasoline by station brand type are shown using the two-stage DID estimator in Gardner et al. (2024). Hypermarts are omitted from Panel B as there are too few hypermarts that exit the sample. Data are sourced from the Oil Price Information Service (OPIS).

In Figure 12, we compare the effects of the entry or exit of a station on nearby stations that share the same store branding, for example, two nearby 7-eleven branded stations regardless of the fuel type they choose to sell. Although we continue to find null effects for exiting stations, we find modestly larger impacts of entrants on incumbent stations with the same brand. In Appendix Figure A6, we report nearly identical results when we instead compare the brand of gasoline sold by the stations.

Figure 12: Effect of Station Entry and Exit by Station Branding



(a) Station Entry



(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in panel A) and exit (in panel B) within 1 mile of an incumbent station on incumbent pricing using the two-stage DID estimator in Gardner et al. (2024) are shown. Results are plotted separately for incumbent stations that share and do not share the same store brand as the entering or exiting station at the time of the event. Data are sourced from the Oil Price Information Service (OPIS).

5.3 Selection of exiting stations

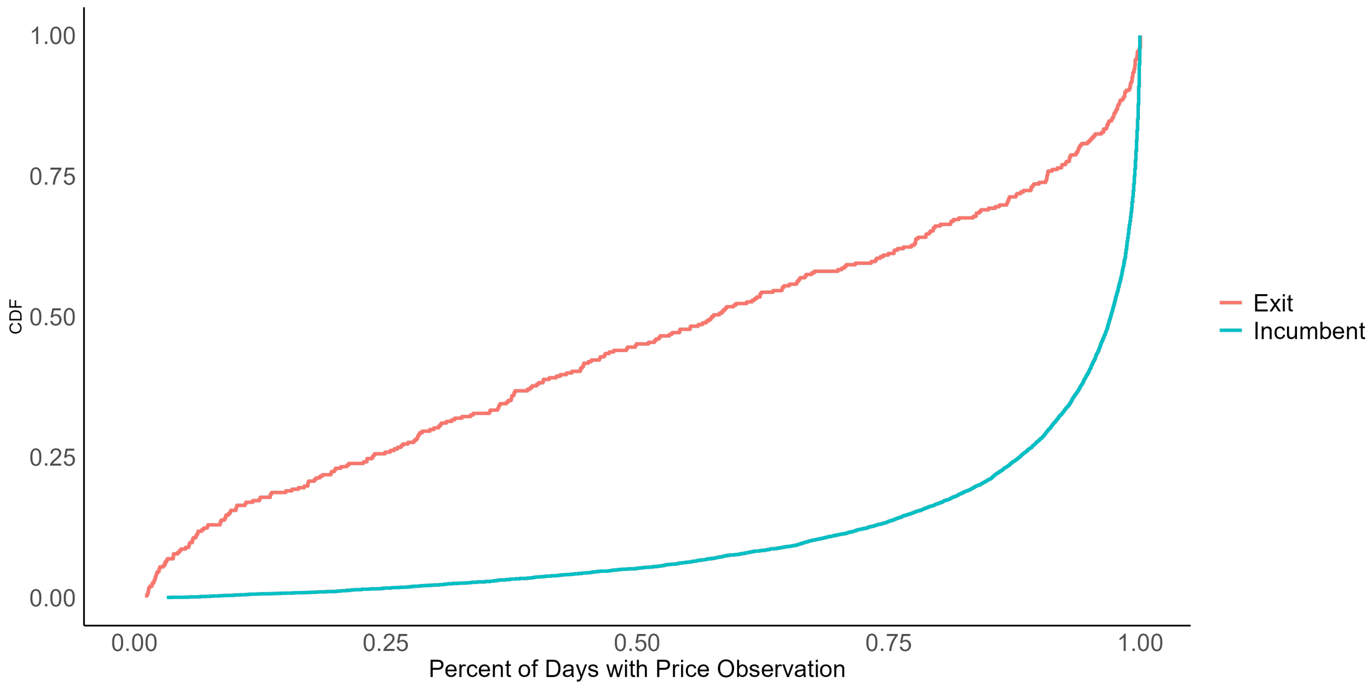
Next, we consider whether selection plays a role in explaining the null effects of exit we observe. If, prior to exit, marginal stations exert little competitive pressure on nearby stations, either because they are not cost-competitive or fail to offer a "product" (inclusive of station attributes) that is attractive to potential customers, their exit might not impact prices at neighboring stations.

We return to the OPIS data to construct a proxy for the "unobserved competitiveness" of exiting stations by examining the frequency with which exiting stations report prices. OPIS collects price data through different streams, two of which, card swipes and consumer-reported prices through GasBuddy, require a customer to use (or observe prices) at a station. To the extent that an exiting station is unattractive to customers, prices may be observed less frequently prior to exit.

As illustrative evidence, we calculate the frequency of reported prices for each station in the OPIS data. Figure A3 graphs the CDF for incumbent stations (in blue) and exiting stations (in red). As noted in Section 2, prices tend to be regularly reported for the vast majority of stations in the OPIS data – the median station in the OPIS data has an observed price on 97% of days. We see this reflected in the CDF for incumbent stations, for which the OPIS data reports prices almost every day. In sharp contrast, we see that price reporting is much less regular for exiting stations.[12] On average OPIS reports a price roughly every other day for the median exiting station, starkly less than the (close to) flawless reporting for the median incumbent station.

[12]As before, when calculating the frequency for exiting stations, we only include in the denominator of the calculation the number of days prior to our observed exit.

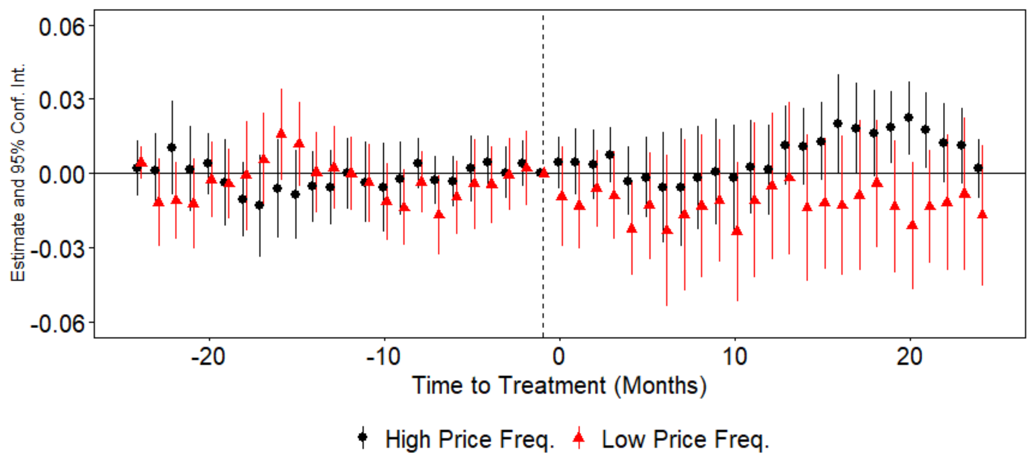
Figure 13: Frequency of reported prices



Notes: The CDF for the percent of potential days a station has a valid price reported in the OPIS data is shown. The percentage is calculated as the number of price observations divided by the number of days spanned between the first and last price observation for a station. Exit stations exit at some point during our sample. Incumbents neither exit nor enter during the sample period.

To examine the role of selection, we split exiting stations based on the frequency with which OPIS reports prices prior to station exit. We then estimate the impact of station exit on nearby station pricing for exiting stations with above-median reporting frequency and below-median reporting frequency. Event study coefficients for the high and low-frequency exiting stations are plotted together in Figure 14. Comparing the estimates for high frequency exiting stations (in black) and low frequency exiting stations (in red), we find suggestive evidence that the exit of a ”high-frequency” station impacts incumbents differently than the exit of a ”low-frequency” station. The point estimate for the impact of the exit of the high-frequency stations is consistently above that of the low-frequency stations consistent with the narrative that these stations are more likely to be relevant competitors in their markets.

Figure 14: Effect of Station Exit, by Reporting Frequency



Notes: Coefficient estimates for the effect of station exit within 1 mile of an incumbent station on incumbent pricing, separated by exiting stations with above and below median reporting frequency prior to exit, are shown. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

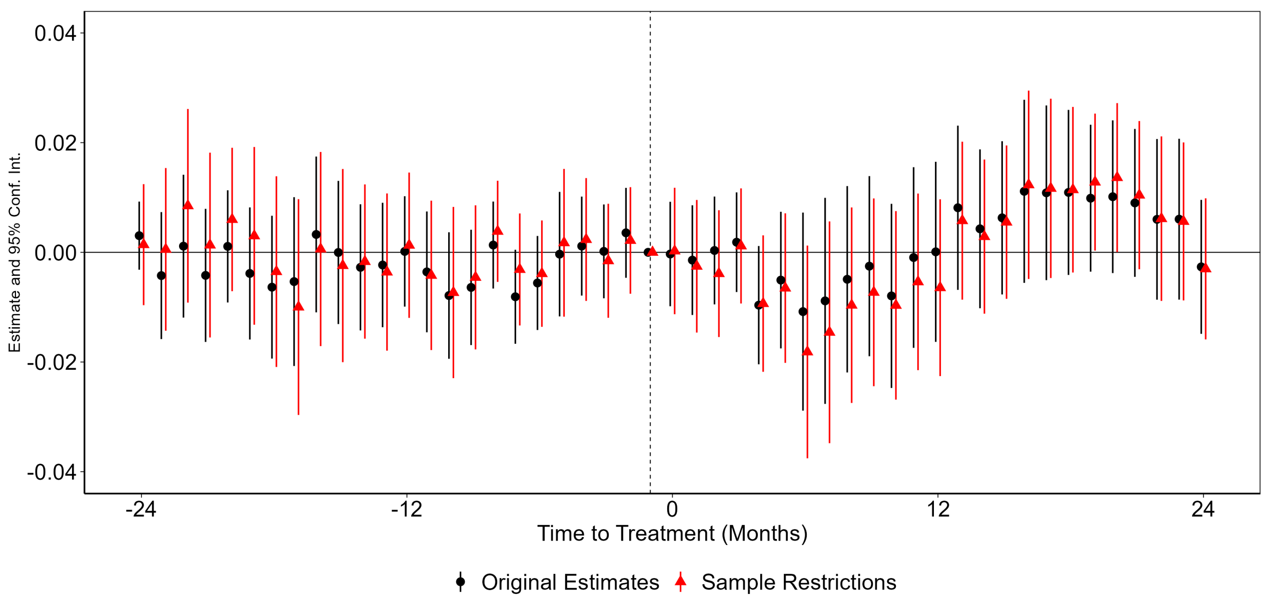
5.4 Misclassification of exiting stations

Finally, we consider whether misclassification of exiting stations might provide an explanation for the null results we observe for exit events. As discussed in the data section, we impute the entry and exit dates of stations from the first and last price observations in the data and classify a station as having exited if we do not observe a price for that station during the last two months of our sample period. If a station has highly sporadic price observations, we could misclassify the first or last observation as an entry or exit, when in reality they had subsequent observations outside the bounds of our sample.

To test this, we make two sample restrictions to increase our confidence that we are identifying actual exit events. First, we no longer consider any station that has less than 25% of potential days with a reported price as an exit. Secondly, we originally did not consider first price observations in the first 3 months or last observations in the last 2 months of the sample to be entries or exits, and instead considered these stations to be active throughout the entire sample. We increase that restriction to no longer consider entries and exits in the first and last 6 months of the sample. As an example, a station whose last price is reported in August 2018 will no longer be considered an exiting station.

In figure 15 we find little difference in our estimates after excluding the exiting stations for which misclassification might be the most prevalent. We continue to find a null effect of exit on incumbent station pricing.

Figure 15: Effect of Station Exit on Incumbent Pricing



Coefficient estimates for the effect of station exit within 1 mile of an incumbent station on incumbent pricing are shown. Original estimates are shown in black. The specification with additional sample restrictions shown in red excludes exiting stations that report a price on less than 25% of days and does not consider stations that exit events in the final 6-months of the sample. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

6 Conclusion

Using daily price data and the timing of the entry of new gas stations and exit of existing gas stations, we estimate the effect of market size changes on the pricing for incumbent stations. The use of high-frequency data and the ability to restrict identification to within-station variation allows the difference-in-differences and event study approaches to account for the endogenous entry and exit decisions of firms. We find that an increase in market size from entry is associated with a statistically significant 2-cent decrease in the price at incumbent stations, reflecting a 5% decrease in average retail markups. This is compared to a precise null effect for the exit of a nearby firm. Both results are robust to various specifications, new estimators that correct for heterogeneous treatment effects and differential treatment timing in the two-way fixed effects specification, and across the various blends of gasoline sold. The results are strongest for the closest entries and dissipate as the market definition broadens. These results are in line with and of similar magnitude to recent studies in other countries (Davis et al. 2023; Fischer et al. 2023).

In contrast, we estimate precise null effects for exit. We note two features that might help to explain the asymmetry between the effects of entry and exit in our setting. First, entry tends to occur in more concentrated markets than exit, where the impact of a change in competition is greater. In addition, we find suggestive evidence that exiting stations differ from incumbent (or entering) stations. Although there are no differences on observable dimensions that explain the null effects for exit, we find differences when separating exiting stations based on the frequency with which prices are reported. Notably, when stations with more frequently reported prices exit, prices rise at incumbent stations. As OPIS reporting relies (as least partially) on transaction data and cloud-sourced price reports, the frequency with which prices are reported serves as a potential proxy for unobservable station characteristics. If infrequently-reporting stations tend to exert less competitive pressure on neighboring firms, their exit might not impact prices.

This paper contributes to the growing body of evidence documenting the competitive effect in retail gasoline markets and offers, to our knowledge, the first causal estimates for the effect of entry and exit of gasoline stations in California. This result is important for policy discussions surrounding market power of retail gas stations. Additionally, this work contributes to the policy discussion surrounding the energy transition, showing that restricting needed expansion of fueling infrastructure could lead to a distortionary price effect while also enriching owners of existing legacy infrastructure.

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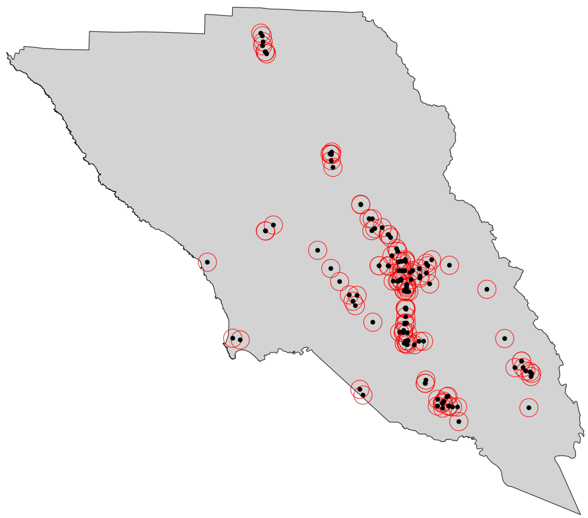
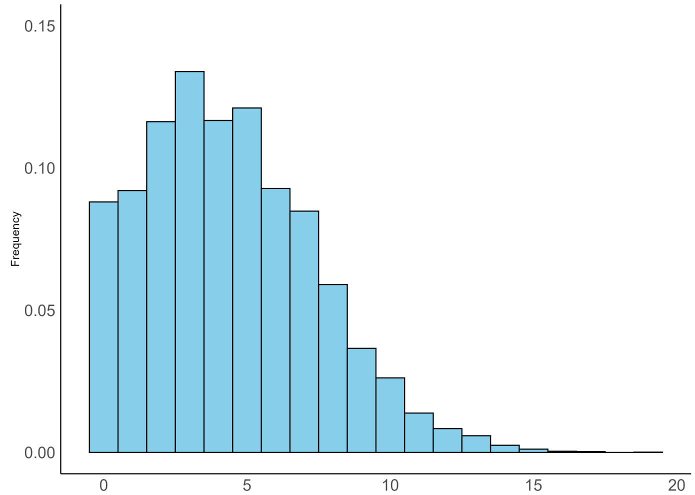
A Appendix: Figures

Appendix Figure A1: Map of Gasoline Stations in Sonoma County: 2014-2018



Notes: The 1 mi. market definition for stations in Sonoma county, California are shown. Data sourced from Oil Price Information Service (OPIS)

Appendix Figure A2: Initial Market Size by Station



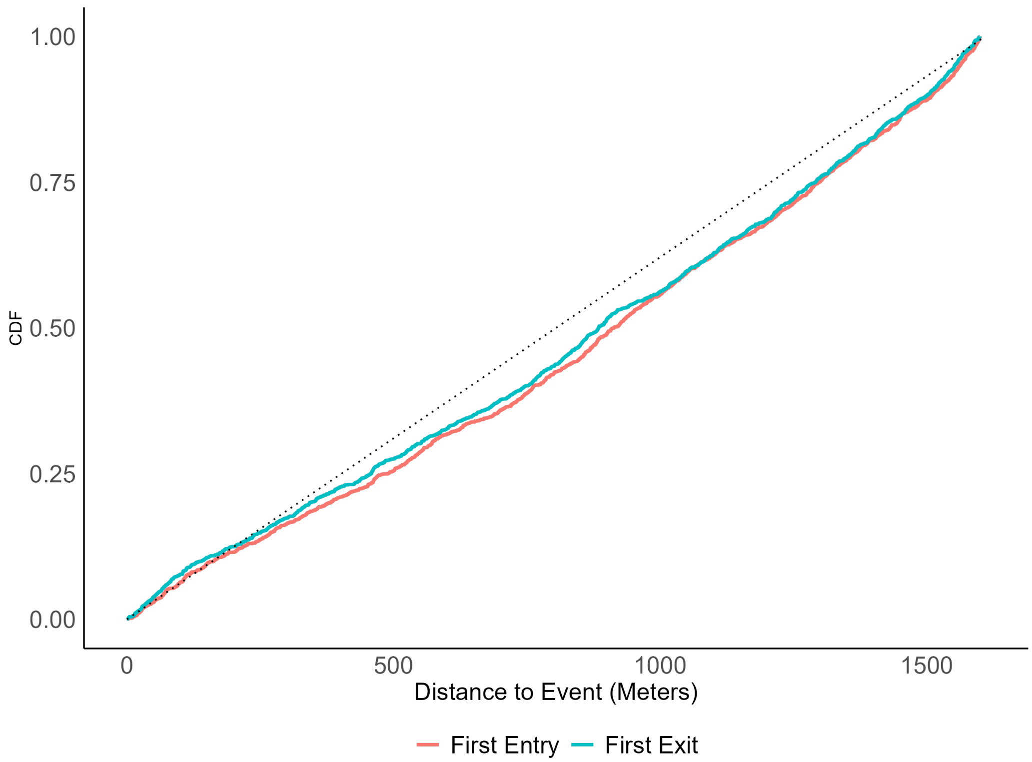
Notes: The number of other stations within 1 mile on the date of the earliest reported price observation is shown.

Appendix Figure A3: Number of Entrants per Incumbent Station by Initial Market Size

		Number of Entrants						
		0	1	2	3	4	5	Total
Initial Market Size	0	783	43	3	1	0	0	830
	1	781	81	12	3	0	0	877
	2	968	124	9	3	0	0	1,104
	3	1,084	163	25	2	0	0	1,274
	4	919	158	19	1	1	0	1,098
	5	970	153	15	2	6	1	1,147
	6	765	120	9	2	1	0	897
	7	636	165	14	2	0	0	817
	8	438	101	13	7	2	0	561
	9	289	54	12	0	0	0	355
	10	207	37	11	0	0	0	255
	11	114	25	2	0	0	0	141
	12	57	16	1	1	0	0	75
	13	48	13	0	0	0	0	61
	14	16	6	1	1	0	0	24
	15	9	3	0	0	0	0	12
	16	3	1	0	0	0	0	4
	17	3	0	0	0	0	0	3
	19	1	0	0	0	0	0	1
	Total	8,091	1,263	146	25	10	1	9,536

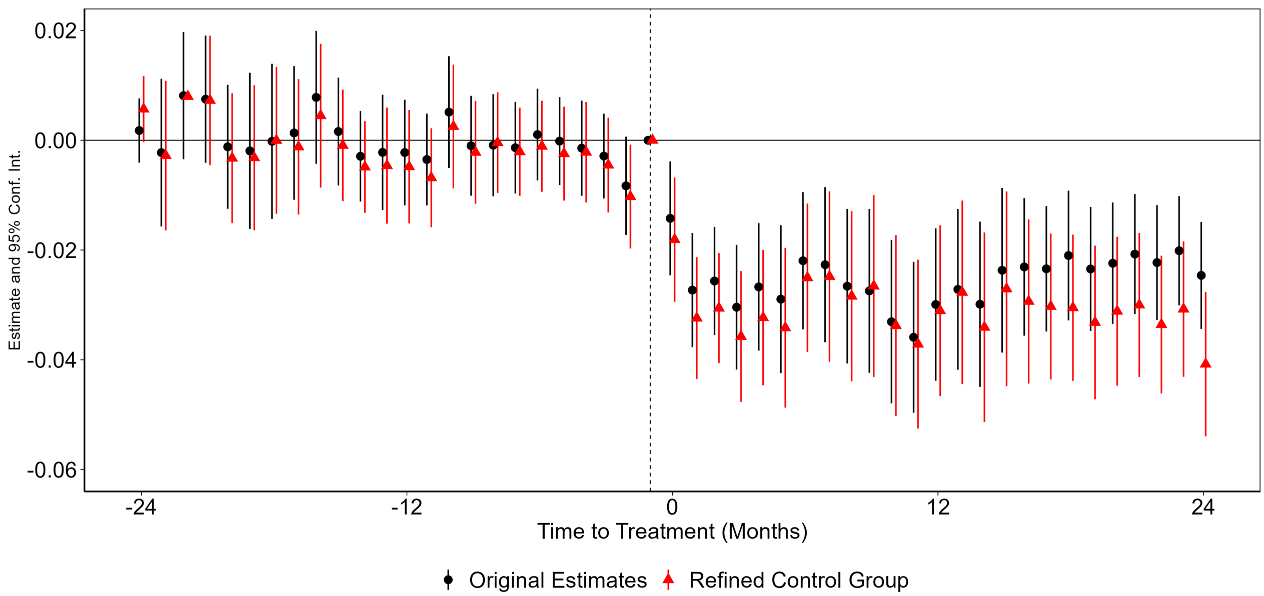
Notes: The number of entrants experienced for each station is shown by the initial market size for the incumbent station.

Appendix Figure A4: Distance from Incumbent Station to Entrant/Exiting Station



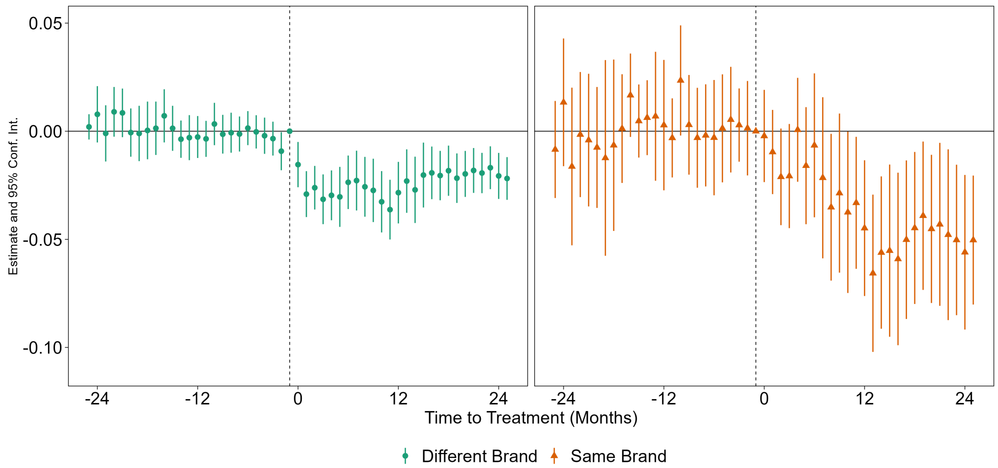
Notes: The figure plots the CDF of the distance from the incumbent station to the identified entrant or exiting station. Data sourced from Oil Price Information Service (OPIS)

Appendix Figure A5: Effect of Entry Within 1 Mi. on Incumbent Pricing Excluding Intermediate Entry Distances

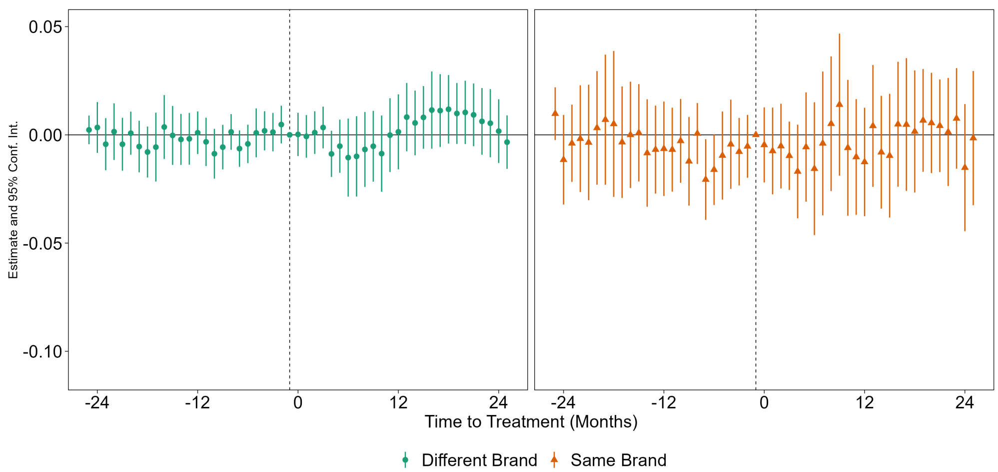


Notes: Coefficient estimates for the effect of station entry within 1 mile of an incumbent station on incumbent pricing are shown. The refined control group removes control stations which experienced an entry within 7 miles during the sample. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

Appendix Figure A6: Effect of Station Entry and Exit by Gas Branding



(a) Station Entry

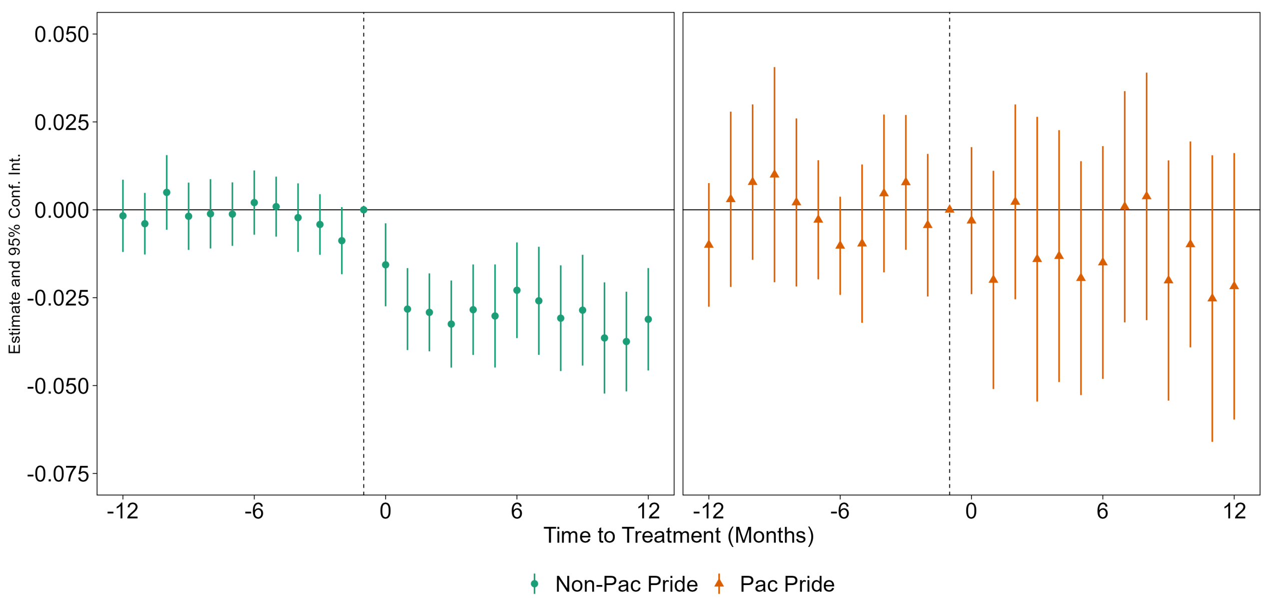


(b) Station Exit

Notes: Coefficient estimates for the effect of station entry (in panel A) and exit (in panel B) within 1 mile of an incumbent station on incumbent pricing are shown. Results are plotted separately for incumbent stations that share and do not share the same gasoline brand as the entering or exiting station at the time of the station closure. Data are sourced from the Oil Price Information Service (OPIS).

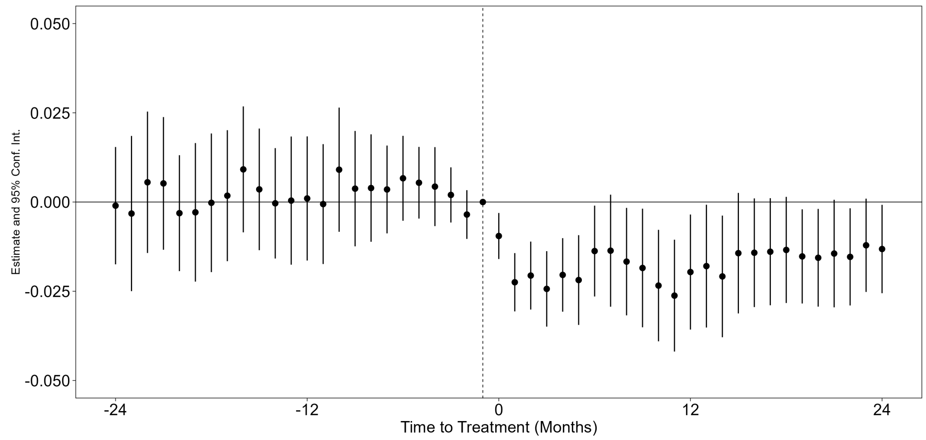
Looking at the store brands in the OPIS data, 38 stations operate under the Pacific Pride USA branding. These are commercially focused gas stations which service mostly fleet vehicles, require a membership and store specific payment card, and are do not have on-site attendants or convenience stores. These events can serve as a falsification test, as nearby entry of a Pacific Pride USA station should have a muted effect on stations which are mostly serving non-fleet vehicles. In Figure A7 we plot the coefficients from regressions for Pacific Pride USA stations and non-Pacific Pride USA stations. We focus on the year before and after entry due to the data further away from the event. The entry of a Pacific Pride USA station is not associated with a spillover on to the pricing of incumbent stations, leading to more precise and a deeper price effect estimate after their removal.

Appendix Figure A7: Effect of Pac Pride Entry Within 1 Mi. on Incumbent Pricing



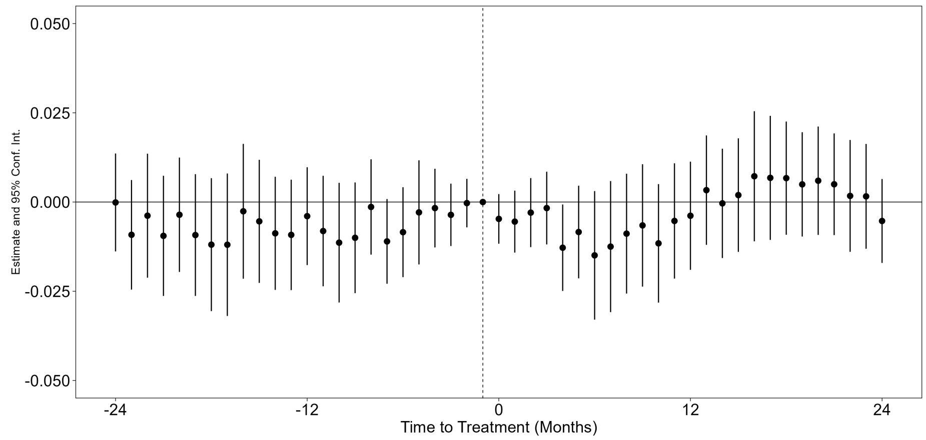
Notes: Coefficient estimates for the effect of station exit within 1 mile of an incumbent station on incumbent pricing for Pacific Pride USA stations and all other stations are shown. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

Appendix Figure A8: Effect of Station Entry on Incumbent Pricing



Notes: Coefficient estimates for the effect of station entry within 1 mile of an incumbent station on incumbent pricing of regular unleaded gasoline are shown. Event time t=-1 is the month prior to the arrival of the entrant. Coefficients are estimated from a linear regression of price on a panel of event time dummy variables, station fixed effects, day-of-sample fixed effects, and city-specific linear time trends. Standard errors are clustered at the city level. Data are sourced from the Oil Price Information Service (OPIS) for all stations in California for 2014-2018.

Appendix Figure A9: Effect of Station Exit on Incumbent Pricing



Notes: Coefficient estimates for the effect of station exit within 1 mile of an incumbent station on incumbent pricing of regular unleaded gasoline are shown. Event time t=-1 is the month prior to the departure event. Coefficients are estimated from a linear regression of price on a panel of event time dummy variables, day-of-sample fixed effects, and city-specific linear time trend Standard errors are clustered at the city level.