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How Full Is the tank? – Insights on Short-run Fuel Price Reactions from German Travel Diary Data

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Nolan Ritter, Christoph M. Schmidt, and Colin Vance[1]

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Abstract

We provide evidence that motorists respond to short-run fluctuations in fuel prices at the gas pump and not on the road. Employing variants of censored panel regression to control for unobserved heterogeneity and censoring of the dependent variable, we find that the fuel price has a large and negative impact on the quantity of fuel purchased, but no significant impact on the subsequent distance driven per day until the next refill. Over the short-run, drivers thus appear to cope with high fuel prices by adjusting fuel purchases with each visit to the filling station, but without altering their daily mileage.

JEL Classification: C33, Q41, R41

Keywords: Panel data, households, driving behavior, tanking behavior, fuel price

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1 Introduction

The estimation of short-run fuel price reactions is of relevance to transportation policy for at least two reasons. First, these estimates provide insights into how motorists cope with increases in fuel prices when longer run behavioral responses, such as purchasing a new car or changing residential location, are effectively precluded. Second, as a lower-bound estimate of the response to higher fuel costs, the short-run elasticity affords policy-makers with a conservative indicator of the likely effectiveness of price-based instruments in influencing driving behavior.

To date, truly short-run analyses – ones that conclusively hold fixed the role of technology and other long term influences – are relatively smaller in number than long-run analyses or those that draw no temporal distinction. Indeed, as Graham and Glaister (2004, p. 271) note, there is no clear consensus of what constitutes the short- or long-run, with the temporal threshold differing across studies. Hughes et al. (2008), for example, derive short-run elasticity estimates from pooled data measured at a monthly frequency, while Goodwin's (1992) and Crandall's (1992) reviews designate the short-run as generally referring to any period of a year or less. Alternatively, in her meta-analysis of fuel elasticities, Espey (1998, p. 288) denotes the short-run based not on time but on the empirical specification, suggesting that models which include some measure of vehicle ownership and fuel efficiency capture the "shortest" short-run by isolating the influence of price and income changes.

Drawing on daily travel survey data from Germany collected over a period of six weeks, the current study contributes to the above literature with an analysis that definitively isolates the short-run impact of fuel prices on tanking and driving behavior. Several features distinguish the analysis. First, we operate with the highest possible frequency of observations by taking into account all fuel purchases with each visit to the filling station as well as the daily distance driven following the visit.

Second, the model includes household fixed-effects, which, together with the tight time interval separating observations, allows us to control for a wide range of unobservable variables that could otherwise bias the estimates. Third, to accommodate constraints imposed by the size of the fuel tank that may prevent motorists from purchasing the desired amount of fuel, our fixed-effects models additionally incorporate the censoring of the dependent variable using a technique proposed by Alan et al. (2011). Fourth, because the six week survey is conducted annually, we are able to conduct confirmatory analysis by running a model for each year between 2002 and 2010 separately. This affords the unique opportunity to assess the stability of our model results by using information from an identical data generation process. Lastly, rather than relying on an average fuel price as calculated over some arbitrary observation interval, we employ the fuel price that the household actually paid and link this to driving behavior directly following the fuel purchase.

Our findings indicate that the fuel price reaction with respect to the quantity tanked is statistically significant but that its magnitude varies considerably across the samples in our observation period, with average reductions in fuel purchases between -0.35 and -0.85 liters per Euro cent of fuel price increase. Conversely, we find that the fuel price does not determine the average distance driven between refills. In interpreting these results, we contrast our estimates with those of Frondel and colleagues (2008; 2009; 2011), who use the same data source but obtain considerably higher estimates of the effect of fuel price on the distance driven.

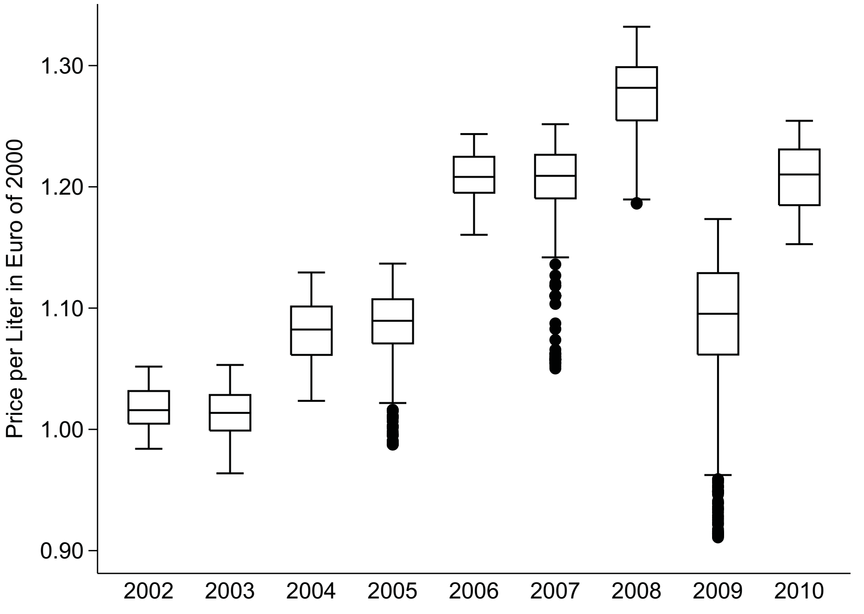
2 Data

This paper uses data taken from the German Mobility Panel (MOP, 2011), an ongoing travel survey that annually collects information on the mobility behavior of a representative sample of German households. We focus here on a subset of this data referred to as the "tank survey", which until 2008 draws a 50% sub-sample of randomly selected car-owning households from the larger MOP. As of 2009, the full sample of car-owning households is surveyed. The tank survey takes place over a roughly six-week period in the spring, during which time respondents record various information upon each visit to the gas station, including the price paid for fuel, the amount of fuel purchased, and the odometer reading. Participating households complete the tank survey upwards of three times over three consecutive years, with exiting households replaced by a new cohort.

Frondel and colleagues sum the distance traveled over the entire six weeks of the survey, and use this sum as the dependent variable in a panel set-up that defines the household as the cross-sectional unit and the year of the survey as the temporal unit. The present analysis takes a different tact. We maintain the household as the cross-sectional unit but structure the temporal dimension of the panel based on the days elapsed between each visit to the gas station within the six week survey period. Distance traveled between each visit is calculated based on the difference in the odometer reading and normalized by dividing by the number of elapsed days. The resulting measure of daily distance traveled serves as one of our dependent variables. The other dependent variable is defined by the percent of the tank filled directly following each visit to the filling station.

Observations are made at the level of a car, therefore households that own more than one car appear more than once in the dataset. Given that the fixed-effects are specified at the household level, this feature upsets the panel structure of the data as the observations are not uniquely identified by the combination of the cross-sectional and temporal units. We consequently estimate the models on two samples of the data. Our main focus is on a sample that is limited to households owning just one car, which comprises roughly 62% of all car owning households in Germany (Ritter and Vance, 2013). We perform a robustness check on a sample that is expanded to include multicar households. For these households, we select the car that

Figure 1: *Fuel Price by Fuel Type in Euros of 2000*



has the highest reported mileage over the survey period. This ensures that each household, whether single or multi-car, appears one time per year in the data.

Two further adjuestments are made to the sample. First, we remove households who reported taking a car vacation during the observation period as such episodes are unlikely to be representative of short-run driving behavior. Second, recognizing that diesel fuel is not only of higher energy content but also considerably cheaper in Germany compared to gasoline, we remove observations on diesel cars.

A separate data set is created for each year between 2002 and 2010. From Figure 1, we see that the real fuel price in Euros of 2000 for gasoline increased steadily between 2002 and 2008, when the gas price peaked at an average of roughly 1.28 Euros/liter. Table 1 presents descriptive statistics for the variables used in the models as well as variables that describe the structure of the data by year. Between 2002 and 2010, the average percentage of the tank filled when pulling into the gas station was relatively constant, varying sporadically between 0.259 and 0.296, while the average percentage of the tank filled following tanking fluctuates between 0.858 and 0.949.

Table 1: *Descriptive Statistics*

Year	**Obs.**	**Real price**	**% of tank filled**		**% of full refills**	**km/day**
		Gasoline	Before Refill	After refill		
2002	2,401	1.016	0.296	0.949	0.814	29.84
		(0.017)	(0.172)	(0.136)	–	(26.43)
2003	2,071	1.013	0.284	0.935	0.814	30.72
		(0.021)	(0.191)	(0.163)	–	(23.95)
2004	2,326	1.080	0.298	0.912	0.731	28.76
		(0.027)	(0.197)	(0.182)	–	(23.25)
2005	2,224	1.088	0.264	0.858	0.676	30.65
		(0.029)	(0.176)	(0.233)	–	(25.18)
2006	2,128	1.206	0.281	0.882	0.707	29.16
		(0.019)	(0.191)	(0.219)	–	(25.14)
2007	2,236	1.208	0.295	0.888	0.700	29.90
		(0.033)	(0.197)	(0.211)	–	(25.03)
2008	1,994	1.283	0.295	0.875	0.673	29.84
		(0.032)	(0.199)	(0.214)	–	(25.52)
2009	4,206	1.102	0.259	0.874	0.707	30.41
		(0.049)	(0.184)	(0.223)	–	(26.05)
2010	4,450	1.208	0.288	0.883	0.706	30.63
		(0.027)	(0.195)	(0.214)	–	(26.87)

The column labeled Full Tank indicates the share of tanks completely filled. km / day is for kilometers driven per day. Obs. is for Observations. Standard deviations in parentheses.

Turning to the penultimate column, there is some evidence for a pattern wherein higher fuel prices are associated with a lower percent of gas station visits that end in a full tank of gas. In 2002, for example, 81% of gas station visits were full refills, a share that declined to a nadir of 67% in the year 2008 when the gas price peaked. The final column shows that despite a 26%

increase in real fuel prices between 2002 and 2008, mileage has remained fairly constant through the years, averaging about 30 kilometers per day.

3 Modeling issues

Our point of departure in econometrically estimating the short-run response to fuel price fluctuations is the specification of two fixed-effects regressions. Model 1 relates the share of the gas tank of car i filled with fuel on visit t to the fuel price at the station and the share of the tank filled with fuel before refilling:

$$\text{share after}_{it} = \alpha + \beta_1 \cdot \text{price}_{it} + \beta_2 \cdot \text{share before}_{it} + \phi_i + \epsilon_{it} \ . \tag{1}$$

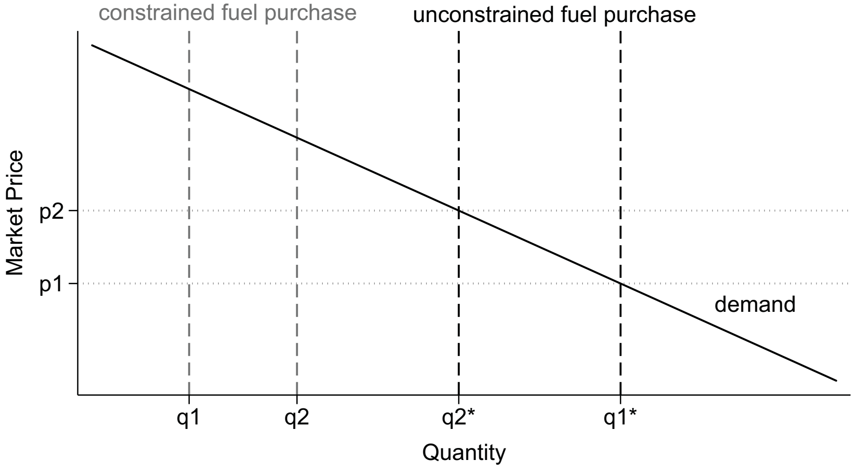
Model 2 relates the daily distance traveled by car i in the interval following a filling station visit t to the fuel price paid at that visit:

$$\text{distance}_{it+1} = \delta + \gamma_1 \cdot \text{price}_{it} + \phi_i + \eta_{it+1} \ . \tag{2}$$

The term ϕ_i represents unobserved household-specific characteristics that affect the outcome variable but do not change over time, while ϵ_{it} and η_{it} are random errors. The key parameters of interest, β_1 and γ_1, measure the effect of the fuel price on the amount tanked and the distance driven.

Whether we can interpret this effect as causal depends critically on our ability to control for the range of confounding factors that determine tanking and driving behavior and that are correlated with the fuel price. One such factor is the proclivity of the motorist to search for cheap fuel; deal-seeking motorists may be more willing to drive extra distance to secure a low fuel price (Yatchew and No, 2001, p. 1706). If so, the omission of this character trait from the model would impart a negative bias on the estimate of γ in a model of distance driven. The virtue of including a time-invariant and household-specific fixed-effect, ϕ_i, in the model is to control for such

Figure 2: *Price-Quantity Diagram for Fuel*



unobservables. The key identifying assumption is that there are no relevant time-varying unobservable variables. Given the tight temporal linkage between the fuel price and the outcome variables, this assumption would seem relatively benign; it is difficult to conceive of relevant omitted factors that vary over the time between gas station visits.

With respect to the question of tanking behavior, a second empirical concern relates to the constraint imposed by the size of the gas tank on motorists' ability to optimize the volume of gas purchased. When the fuel price is low, for example, it is likely that the motorist will be prevented from purchasing the quantity of fuel desired because of the capacity constraint of her fuel tank. Given that this constraint is binding, we cannot rule out the possibility of a biased coefficient on the fuel price. Moreover, it is not possible to bind the direction of the bias from above or below.

Figure 2, which is a stylized price-quantity diagram for fuel, illustrates an example of this problem. The vertical axis shows two hypothetical prices that prevail for two distinct trips to the gas station at time 1 and time 2. The horizontal axis shows the corresponding maximum amount of fuel that can be purchased given the size of the tank and given the amount of gas already in it when the driver arrives at the station, at q1 and q2. Note that with both visits, the driver is unable to purchase the desired amount of fuel, indicated by the intersection of the price level with the demand curve (q1*, q2*). Under this circumstance, when the price increases from p1 to p2 between the two visits, the driver actually increases the amount of fuel purchased, from q1 to q2. This seemingly perverse demand response emerges because the capacity constraint on the amount of fuel that can be purchased binds for both visits, though it is more relaxed on the second. Of course, other examples could be conceived that illustrate the expected positive demand response to fuel price decreases or a negative response to price increases. But even in these cases, the response may either be relatively muted or relatively strong depending on where the constraint lies with each visit to the gas station and on whether it binds.

The crux of the problem is that we are dealing with a censored dependent variable having an upper bound of 1 and a lower bound of 0. Let the model under consideration be

$$y_{it}* = \boldsymbol{x}_{it} \cdot \boldsymbol{\beta} + \phi_i + \epsilon_{it} \ , \tag{3}$$

where $\boldsymbol{x}$ is a vector of control variables with a corresponding vector of parameters $\boldsymbol{\beta}$ and an error term ϵ. Rather than observing $y*$, we observe

$$y_{it} = \begin{cases} L & \text{if } y_{it}* < L \\ y_{it}* & \text{if } L \leq y_{it}* \leq U \\ U & \text{if } y_{it}* > U \ , \end{cases} \tag{4}$$

where L (U) indicates the lower (upper) constraint (Alan et al., 2011).

In a panel setting, a standard econometric solution to this problem is the random effects tobit model. This model has a major drawback, however, in that it requires us to assume that ϕ_i and the explanatory variables are uncorrelated. As an alternative, we use a method recently proposed by Alan et al. (2011) that is particularly useful when the dependent variable is a fraction, as in the present case. This model builds on earlier work by Honoré (1992), who developed a semi-parametric estimator for the fixed-effects tobit model, one based on the construction of moment conditions for panel data models with one-sided truncation or censoring. The estimator used here, also based on the construction of moment conditions, generalizes that approach to the case of two-sided truncation or censoring. Referred to as a two-sided censoring model (TSCM), it is thus ideally suited to the present application, where we have two-sided censoring at 0 and 1 given by the share of the tank filled with fuel.

4 Results

We begin by presenting a model of the amount tanked using a standard fixed-effects estimator. Thereafter, we control for the effects of censoring by presenting the TSCM of Alan et al. (2011). Finally, we complete the presentation with a fixed-effects model of distance traveled.

4.1 Tanking behavior

Table 2 presents coefficient estimates by year from a standard fixed-effects model of the amount tanked focussing on single-car households. With the exception of the years 2002 and 2004, the results indicate that the fuel price has a statistically significant and negative impact on the amount of fuel purchased during a filling station visit. Moreover, the magnitude of the estimates varies considerably.

The smallest significant estimate is seen for the year 2009, when a 1 cent increase in the fuel price is associated with a decrease in the amount of fuel purchased on the order of 0.6 percentage points of tank volume (0.0062). With an average tank volume of 55 liters in the sample, this amounts to a decrease of about 0.34 liters. This is contrasted by a considerably higher estimate for the year 2010, which, at 0.0159, is about two and a half times the magnitude of the estimate for 2009. Likewise, the estimates of the effect of the share of fuel in the tank prior to filling up (*share before*) is highly variable. In 2008, a 1 percentage point increase in the share of the tank filled prior to the refill increased the share of the tank filled after the refill by 0.31, which is over five times the magnitude of the corresponding estimate for 2003.

Table 3 presents estimates from the sample that includes both single- and multi-car households. Overall, the same pattern emerges. The smallest significant estimate is 0.0050 for the year 2002, which, unlike above, is now statistically significant. The largest estimate is 0.0181 for the year 2010, which is over three times higher in magnitude. The pattern of estimates on the variable *share before* is also roughly the same as in the model limited to single-car households.

We additionally estimated alternative specifications that included other covariates. For example, we tested whether the time of month is an important correlate under the hypothesis that at later dates following payday there would be less disposable income for households to spend on fuel. We also explored the influence of the local average weekly temperature and precipitation, using a geographical information system (GIS) to merge these variables with the data. As none of these variables were found to be statistically significant, they were left out of the final specifications.

Table 2: *Standard Fixed-Effects in Single-Car Households*

		2002 N=788	**2003** N=753	**2004** N=651	**2005** N=655
real price		−0.0056	−0.0079**	−0.0027	−0.0083**
		(0.0028)	(0.0024)	(0.0022)	(0.0027)
share before		0.1811**	0.0578*	0.1781**	0.2461**
		(0.0300)	(0.0266)	(0.0364)	(0.0408)
	2006 N=635	**2007** N=699	**2008** N=658	**2009** N=1,340	**2010** N=1,343
real price	−0.0112**	−0.0121**	−0.0073**	−0.0062**	−0.0159**
	(0.0034)	(0.0028)	(0.0020)	(0.0011)	(0.0019)
share before	0.1657**	0.2256**	0.3107**	0.2642**	0.2683**
	(0.0410)	(0.0478)	(0.0447)	(0.0398)	(0.0307)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

Table 3: *Standard Fixed-Effects for Most Used Car*

		2002 N=1,152	**2003** N=987	**2004** N=992	**2005** N=973
real price		−0.0050*	−0.0061**	−0.0032	−0.0081**
		(0.0023)	(0.0021)	(0.0018)	(0.0022)
share before		0.1510**	0.0599*	0.1874**	0.2138**
		(0.0251)	(0.0244)	(0.0305)	(0.0360)
	2006 N=911	**2007** N=944	**2008** N=818	**2009** N=1,760	**2010** N=1,796
real price	−0.0128**	−0.0117**	−0.0079**	−0.0066**	−0.0181**
	(0.0030)	(0.0023)	(0.0018)	(0.0009)	(0.0017)
share before	0.1914**	0.2088**	0.2939**	0.2705**	0.2457**
	(0.0376)	(0.0388)	(0.0393)	(0.0334)	(0.0280)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

To explore the extent to which the estimates in Tables 2 and 3 may be biased by failing to take into account the censoring of the dependent variable, Tables 4 and 5 present the results from the two-sided censoring model. Comparison is facilitated by presenting the marginal effects from these models, which, following Alan et al. (2011), are obtained by multiplying the coefficient estimate by the share of uncensored observations. The coefficient estimates are presented in the appendix.

With respect to the effect of price, the largest discrepancy is seen for the year 2002, when the coefficient in the TSCM for single-car households is roughly 1.7 times higher and statistically significant compared to that of the corresponding fixed-effects model. Otherwise, for most years, the magnitude of the price coefficients in the two sets of estimates are relatively similar. By contrast, the magnitude of the estimates on *share before* are uniformly higher in the TSCM, in some years considerably so. In 2006, the estimate of the TSCM is nearly double that of the fixed-effects model. As indicated in Table 5, including multi-car households in the sample has little bearing on the estimates obtained from the TSCM.

Perhaps the most intriguing aspect of the estimates - irrespective of the estimation method or the household's endowment of cars - is their rather large variation over different years. While it is not immediately evident what accounts for this, one explanation may be differing degrees in the level of fuel price variability across the years. It is conceivable, for example, that motorists would display a higher level of price sensitivity during periods of higher price variability. We plotted the coefficient estimates against a measure of the daily variance in prices for each six-week period in order to glean anecdotal support for this explanation, but no discernible pattern emerged. Irrespective of the source of the variation in the estimates, they illustrate that inferences drawn from a single year of data may mask substantial inter-year heterogeneity.

Table 4: *Marginal Effects for the Two-Sided Censoring Model with Individual Specific Effects for Single-car Households*

		2002 N=788	**2003** N=753	**2004** N=651	**2005** N=655
real price		−0.0097** (0.0037)	−0.0093** (0.0029)	−0.0026 (0.0023)	−0.0076** (0.0024)
share before		0.2511** (0.0399)	0.1101* (0.0504)	0.2704** (0.0479)	0.3185** (0.0399)
	2006 N=635	**2007** N=699	**2008** N=658	**2009** N=1,340	**2010** N=1,343
real price	−0.0108** (0.0035)	−0.0134** (0.0037)	−0.0082** (0.0020)	−0.0068** (0.0012)	0.0150** (0.0023)
share before	0.3103** (0.0566)	0.3168** (0.0599)	0.3735** (0.0480)	0.3751** (0.0444)	0.4393** (0.0386)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

Table 5: *Marginal Effects for the Two-Sided Censoring Model with Individual Specific Effects for Most Used Car*

		2002 N=1,152	**2003** N=987	**2004** N=992	**2005** N=973
real price		−0.0090** (0.0033)	−0.0084** (0.0026)	−0.0037 (0.0021)	−0.0080** (0.0022)
share before		0.2185** (0.0363)	0.1258** (0.0469)	0.2727** (0.0426)	0.3071** (0.0395)
	2006 N=911	**2007** N=944	**2008** N=818	**2009** N=1,760	**2010** N=1,796
real price	−0.0124** (0.0035)	−0.0120** (0.0031)	−0.0083** (0.0019)	−0.0068** (0.0010)	−0.0168** (0.0021)
share before	0.3469** (0.0472)	0.3103** (0.0521)	0.3635** (0.0491)	0.3787** (0.0375)	0.3865** (0.0327)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

4.2 Driving behavior

To assess the impact of fuel prices on driving behavior, Tables 6 and 7 present estimates from a fixed-effects model relating the fuel price to the subsequent daily distance driven following tanking. Table 6 presents estimates from the single-car sample while Table 7 presents estimates from the sample that includes multi-car households. Contrasting with the high degree of price responsiveness displayed at the pump, none of the estimates from these models are statistically significant. Motorists apparently do not adjust driving behavior over such a short time interval in response to changes in fuel prices. Indeed, the high sensitivity revealed at the pump may in part reflect an adaptation mechanism that allows motorists to maintain a steady level of distance driven in the face of price fluctuations.

The absence of significant price effects in Tables 6 and 7 also contrasts with the work of Frondel and colleagues (2008; 2009; 2011) using the same data set. In a series of studies employing a diverse suite of estimators – including panel techniques, sample selection models, and quantile regression – these authors obtain elasticity estimates on the order of -0.6%. As noted above, rather than modeling the distance driven between gas station visits, they measure the total distance driven over the six-week survey period, which is recorded over each of three consecutive years for every household. The temporal dimension of their analyses is thus measured as a year. Evidently, over this longer interval, motorists display higher adaptability to high fuel prices, resulting in less driving.

Table 6: *Driving Distance Between Refills (Standard Fixed-Effects) for Single-Car Households*

		2002	2003	2004	2005
		N=631	N=609	N=539	N=568
real price		0.0984	2.7221	0.7584	−0.2318
		(2.6907)	(1.9979)	(1.1192)	(1.3625)
	2006	**2007**	**2008**	**2009**	**2010**
	N=531	N=593	N=543	N=1,152	N=1,089
real price	−0.4463	−1.7825	1.5231	−0.1730	−1.3493
	(1.6171)	(1.2596)	(0.9223)	(0.4686)	(0.8915)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

Table 7: *Driving Distance Between Refills (Standard Fixed-Effects) for Most Used Car*

		2002	2003	2004	2005
		N=934	N=794	N=818	N=844
real price		0.9809	0.7482	0.2082	0.4787
		(2.3811)	(1.8251)	(1.0600)	(1.0264)
	2006	**2007**	**2008**	**2009**	**2010**
	N=759	N=798	N=679	N=1,515	N=1,477
real price	−1.0085	−1.0521	1.1725	−0.1475	−1.4431
	(1.2804)	(1.0384)	(0.8936)	(0.3983)	(0.7963)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

5 Conclusion

This article employs fixed-effects panel regression techniques to explore the influence of fuel prices on tanking and driving behavior over the very short-run, with the temporal dimension defined as the interval between visits to the filling station over a six week survey period. Beyond estimating a standard fixed-effects model, we estimate a model of tanking behavior that uses a censored regression technique developed by Alan et al. (2011) to address the constraint that households can only buy as much fuel at the pump as the free volume of their tank allows. Taking advantage of the annual availability of the data, we perform confirmatory analysis by repeating the estimation exercise for each year between 2002 to 2010. Our estimates suggest that motorists are very sensitive to fuel prices at the pump, and that the degree of this sensitivity varies considerably across survey years, with the coefficients varying by upwards of 2-fold. Conversely, our estimates of mileage suggest that fuel prices do not bear on the daily distance driven between visits to the filling station; the coefficient estimates from these models are all statistically insignificant.

Given the structure of our data, we conclude that in the absolute short-run, when everything is fixed including the endowment with cars, habits, living environment, and transport alternatives, households react to fuel price fluctuations by adjusting the amount tanked but not the amount driven. This finding contrasts with that of Frondel and colleagues, who also conduct what may be deemed as short-run analyses with the same data source, albeit with the temporal dimension defined over a year. That these authors obtain statistically significant fuel price elastcities with respect to driving of about -0.6 highlights that the fuzzy distinctions of what constitutes the short-run can have a fundamental bearing on the conclusions drawn from the analysis.

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6 Appendix

Table 8: *Panel Censored Regression with Two-Sided Censoring and Individual Specific Effects*

		2002 N=788	**2003** N=753	**2004** N=651	**2005** N=655
real price		−0.0353** (0.0136)	−0.0356** (0.0111)	−0.0082 (0.0073)	−0.0197** (0.0064)
share before		0.9117** (0.1448)	0.4231* (0.1937)	0.8628** (0.1529)	0.8310** (0.1042)
	2006 N=635	**2007** N=699	**2008** N=658	**2009** N=1,340	**2010** N=1,343
real price	−0.0268** (0.0088)	−0.0319** (0.0089)	−0.0204** (0.0050)	−0.0162** (0.0028)	−0.0399** (0.0061)
share before	0.7698** (0.1404)	0.7558** (0.1429)	0.9310** (0.1196)	0.8959** (0.1061)	1.1660** (0.1024)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.

Table 9: *Panel Censored Regression with Two-Sided Censoring and Individual Specific Effects for Most Used Car*

		2002 N=1,152	**2003** N=987	**2004** N=992	**2005** N=973
real price		-0.0348** (0.0128)	-0.0325** (0.0102)	-0.0117 (0.0065)	-0.0206** (0.0056)
share before		0.8448** (0.1403)	0.4868** (0.1815)	0.8561** (0.1336)	0.7926** (0.1019)
	2006 N=911	**2007** N=944	**2008** N=818	**2009** N=1,760	**2010** N=1,796
real price	-0.0313** (0.0089)	-0.0299** (0.0077)	-0.0211** (0.0050)	-0.0164** (0.0024)	-0.0436** (0.0054)
share before	0.8755** (0.1191)	0.7769** (0.1304)	0.9264** (0.1251)	0.9167** (0.0908)	1.0017** (0.0846)

** and * indicate significance at the 1% and 5% level. Standard errors in parentheses.