THE PRICE OF GASOLINE AND THE DEMAND FOR FUEL EFFICIENCY: EVIDENCE FROM MONTHLY NEW VEHICLES SALES DATA *

Thomas Klier
Federal Reserve Bank of Chicago
tklier@frbchi.org

Joshua Linn
Department of Economics
University of Illinois at Chicago
jlinn@uic.edu

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ABSTRACT

This paper uses a unique data set of monthly new vehicle sales by detailed model from 1970-2007, and implements a new identification strategy to estimate the effect of gasoline prices on new vehicle demand. We control for unobserved vehicle and consumer characteristics by using within model-year changes in gasoline prices and vehicle sales. We find a significant demand response, as nearly half of the decline in market share of U.S. manufacturers from 2002-2007 was due to the increase in gasoline prices. On the other hand, a gasoline tax increase would have a modest effect on average fuel efficiency.

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1 Introduction

The rapid rise in the price of gasoline from just over $1 at the beginning of 2002 to over $4 by mid 2008 has renewed interest in the relationship between the price of gasoline and the demand for fuel efficient vehicles in the U.S. market. Recent research on oil prices and economic activity suggests that because of improved energy efficiency, the U.S. economy as a whole is currently much less sensitive to oil prices than it was prior to the 1970s oil shocks (Hooker, 1996 and Linn, 2008). In contrast, recent events suggest that U.S. automakers may remain quite sensitive to oil and gasoline prices.

Many industry analysts and the popular press have noted the large decrease in sales for U.S. automakers and large SUVs over the past six years, and have widely attributed some of the changes to the coinciding increase in the price of gasoline. Figure 1 depicts these trends, showing a 20 percent decrease in the market share of U.S. firms between 2002 and 2007, which represents several billion dollars per year in lost profits. The figure also suggests that some of the decrease in sales by U.S. firms can be attributed to changes in the SUV segment of the market. At the beginning of this time period, U.S. firms accounted for 80 percent of sales of large SUVs (with a mean fuel efficiency of 16.6 miles per gallon), but only 37 percent of sales of smaller SUVs (commonly called crossover utility vehicles, with a mean fuel efficiency of 22.2 miles per gallon). The market share of small SUVs has increased at the expense of the market share of large SUVs, and at the same time the real price of gasoline has nearly doubled. Although the contemporaneous correlations between the price of gasoline and market shares are suggestive, it must be recalled that fuel costs are only a fraction of the total cost of the vehicle, and that many

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1 The market share of U.S. automakers had been declining since the mid 1960s. After stabilizing for most of the 1980s and the early 1990s, it started declining again in 1997.
other factors could explain these trends. In fact, despite the attention given to gasoline prices by the press, there has not been a rigorous analysis of the extent two which increases in gasoline prices adversely affect U.S. automakers.

The effect of the price of gasoline on new vehicle demand is also central to the ongoing policy debate over reducing gasoline consumption, which has emanated from concerns about global warming and reducing oil imports. In 2007, Congress passed legislation that increased the CAFE standard by about 40 percent, to be effective by the year 2020. During the Congressional debate over the bill, several members of Congress proposed an increase in the gasoline tax or the introduction of a carbon tax as alternative means of reducing gasoline consumption. The welfare effects of the CAFE standard and the gasoline tax, as well as the optimal gasoline tax (Parry and Small, 2005), depend partly on the effect of the price of gasoline on the demand for fuel efficient vehicles (Austin and Dinan, 2005).

Despite the large empirical literature on the effect of the price of gasoline on new vehicle demand, the cause of the changes in market shares in Figure 1 and the effect of the gasoline tax on average fuel efficiency remain as open questions. Previous studies focus on the effect of gasoline prices on average fuel efficiency across the entire market, and thus do not shed light on the causes of the changes in market shares depicted in Figure 1. Furthermore, as discussed in more detail below, these studies do not control for both unobserved consumer and vehicle characteristics, which may bias the estimated effect of the gasoline tax on average fuel efficiency.

In this paper, we estimate the effect of the price of gasoline on the demand for fuel efficient vehicles. We use a unique data set of monthly sales by vehicle model that spans nearly 40 years.

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2 See, for example, hearings at the House Energy and Commerce committee on March 14, 2007.
3 During the summer of 2007, Congressman Dingell (MI) proposed a 50 cent per gallon gasoline tax increase (Timiraos, 2007)
The disaggregated and high-frequency data allow for a simple and transparent empirical strategy that controls for unobserved consumer and vehicle characteristics. We find that gasoline prices significantly affect the new vehicles market, as the recent price increase explains nearly half of the decrease in market share of U.S. firms. On the other hand, a one dollar increase in the price of gasoline would modestly increase fuel efficiency, by 0.5-1 miles per gallon, which is less than some other recent estimates.

We now discuss the analysis and previous literature in more detail. The empirical strategy addresses the limitations of the previous literature on the effect of the price of gasoline on new vehicle demand. First, previous studies have not accounted for the potential correlation between the price of gasoline and unobserved consumer and vehicle characteristics. Many studies rely primarily on cross-sectional variation in gasoline prices (e.g., Goldberg, 1998, West, 2004 and Bento et al., 2006), and thus depend on the questionable assumption that prices are uncorrelated with unobserved consumer preferences (Chouinard and Perloff, 2007); for example, it is assumed that environmentalists are no more likely to live in states with high gasoline taxes. The failure to account for unobservables is particularly problematic because the direction of the resulting bias cannot be determined from economic theory; unobserved vehicle and consumer characteristics may be positively or negatively correlated with the price of gasoline. An additional limitation of the cross-sectional analysis is that the relationship between the price of gasoline and the demand for fuel efficient vehicles may have changed over time. Recent estimates of the elasticity of gasoline consumption to the price of gasoline (e.g., Hughes et al., 2008) suggest that the

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4 An earlier empirical literature estimated the effect of the price of gasoline on new vehicle demand in the 1970s (see Tardiff, 1980, for a summary). These studies have similar limitations to the more recent ones, however. For example, Boyd and Mellman (1980) find a large effect on demand using cross-sectional variation in vehicle characteristics and prices, but the study does not control for unobserved characteristics.
elasticity has decreased in magnitude. However, it is not known whether the effect of the price of gasoline on the demand for fuel efficient vehicles has changed as well.

Furthermore, time series analysis of the price of gasoline on new vehicle demand only partially addresses these issues (e.g., Small and Van Dender, 2007), because these studies use market-level fuel economy measures and do not control for unobserved vehicle characteristics – for example, weight and power are highly correlated with fuel efficiency. Finally, many earlier studies analyze choices among broad vehicle classes, such as small and large cars. There is substantial heterogeneity within classes, however, which implies that the full response may be greater than these studies indicate.

This paper makes several advances beyond the existing literature. First, we use a unique data set and empirical strategy. The data include monthly national sales by detailed vehicle model from 1970-2007, including a new set of 1970s sales data compiled from print sources. As Berry, Levinsohn and Pakes (1995) argue, when using market-level data it is necessary to account for the potential correlation between vehicle prices, sales and unobserved vehicle and consumer characteristics. The monthly frequency of the sales data allows for a simple linear estimating equation that controls for these unobserved variables. The empirical strategy derives from the details of automobile production, specifically, that unobserved vehicle characteristics do not vary over the model-year, and the fact that consumer tastes are likely to be slow-moving and uncorrelated with monthly variation of the price of gasoline. The empirical specification exploits within model-year changes in the monthly price of gasoline and vehicle sales, while controlling for unobserved characteristics.5 As discussed in more detail below, the potential drawback of this

5 Similarly to this paper, Busse et al. (2005) use within model-year changes in manufacturing incentives to estimate the effect of incentives on transaction prices. The authors use a regression discontinuity design to address the potential endogeneity of the incentives. In contrast, we assume that the price of gasoline is exogenous to these incentives, and discuss this assumption further in Sections 5 and 6.
approach is that we estimate the monthly effect of the price of gasoline on sales rather than a long-run effect, which is important for understanding the causes of recent changes in market shares as well as the effect of raising the gasoline tax. We investigate the long-run effect of the price of gasoline on vehicle sales by including lags and controlling for demand shocks over the previous year. We find that the short run and long run quantity effects are similar in magnitude, validating our analysis and the interpretation of the empirical results.

The second contribution of this paper is that the sample period and unit of analysis allow us to investigate a number of questions about gasoline prices and consumer demand that have not been the focus of previous work. First, the model level data allow us to investigate the causes of recent market trends shown in Figure 1. Second, we use the 37-year panel to consider whether the relationship between gasoline prices and new vehicle demand has changed over time, or if there are asymmetric or lagged demand responses.

The main results are reported in Section 5. The increase in the price of gasoline between 2002 and 2007 explains about 25 percent of the shift from large SUVs to small SUVs and 40 percent of the decrease in the market share of U.S. manufacturers. Thus, the results indicate that the price of gasoline has a significant effect on the new vehicles market, but the effect is smaller than is often suggested. We estimate that the elasticity of average new vehicle fuel efficiency with respect to the price of gasoline is about 0.12, which is about half of the elasticity reported in Austin and Dinan (2005). The elasticity implies that a one dollar increase in the price of gasoline raises average fuel efficiency by 0.5-1 miles per gallon.

In comparison to previous studies, we investigate the importance of functional form assumptions and show that the results are robust to using a number of different estimation models. We also find that the demand response to the price of gasoline is greater when gasoline
prices are high. Finally, we find limited evidence of a lagged response to changes in the price of gasoline and provide evidence in support of the assumption that the price of gasoline is exogenous to consumer preferences.

2 Conceptual Framework: The Price of Gasoline and New Vehicle Demand

This section describes a model of consumer demand for new vehicles to illustrate the relationship between the price of gasoline and equilibrium sales. We use a discrete choice model that is similar to Berry, Levinsohn and Pakes (BLP). The market for new automobiles contains $J$ models, or varieties, with each model indexed by $j = 1,\ldots, J$. The market spans one year, with months indexed by $t$. Individual $i$ derives utility $U_{ijt}$ by purchasing model $j$ in month $t$ according to:

$$U_{ijt} = \alpha(p_j + m_j + f_{jt}) + X_j \beta + \xi_j + \epsilon_{ijt}.$$  

(1)

The purchase price of the model is $p_j$, the expected maintenance costs are $m_j$ and the expected fuel costs are $f_{jt}$. These variables reduce available income and therefore utility, so $\alpha$ is negative. The vector $X_j$ consists of attributes of the model that can be observed by the individual and econometrician, such as engine size. The last two variables in equation (1) are the mean utility from unobserved model attributes, $\xi_j$, and an individual- and model-specific error term. In equation (1), $\alpha$ and $\beta$ are constant across models, which implies that the purchase price, operating costs and attributes in the vector $X_j$ affect the utility of all consumers equally. This assumption is stronger than the random coefficients specification of BLP, but will be relaxed below.

A central feature of equation (1) is that the vehicle’s price, expected maintenance costs, as well as observed and unobserved characteristics, do not vary over the model-year and thus do not
have time subscripts. In contrast, expected fuel costs, which depend on the expected price of gasoline, may vary by month. This framework is consistent with the typical production process of new vehicles. For most vehicle lines, production begins in July or August after a brief, one- or two-week, shutdown period. During that period, the manufacturer may change characteristics such as engine size. In practice, changes across model-years range from very minor to a complete overhaul. Once production begins, however, the features of a vehicle are constant over the model-year; a 2005 Honda Civic purchased in September of 2004 is the same as a 2005 Civic purchased in May of 2005. To simplify the notation we define a model-specific intercept that does not vary over time: \( \phi_j \equiv \alpha p_j + \alpha m_j + X_j \beta + \xi_j + \varepsilon_j \). The intercept absorbs vehicle price, maintenance costs, observed characteristics and mean unobserved characteristics, but does not include fuel costs.

We characterize the market-level demand for each model in a straightforward manner. Each consumer purchases one of the \( j \) models or the outside good, which we take to be a used vehicle. Making the standard extreme value assumption for the residual error term yields the following expression for the difference between the log market share of model \( j \), \( s_{j_t} \), and the log market share of the outside good \( s_{0_t} \):

\[
\ln s_{j_t} - \ln s_{0_t} = \alpha f_{j_t} + \phi_j .
\] (2)

Equation (2) is the standard aggregate logit equation with fixed effects, in which the market share of model \( j \) depends negatively on expected fuel costs.

3 Estimating the Effect of Gasoline Prices on New Vehicle Demand

This section describes the empirical strategy for estimating the effect of the price of gasoline on the demand for fuel efficient vehicles. Previous work has used two approaches based on equation...
(1). The first estimates a nested logit model, which requires assumptions about substitution patterns across models and the exogeneity of the price of gasoline (e.g., West, 2004, and McManus 2005). Alternatively, BLP uses market-level data and relies on a set of price instruments as well as assumptions about the distributions of parameters and unobserved characteristics. Our empirical strategy is an alternative to BLP. We use monthly sales data and exploit the fact that vehicle characteristics do not change within the model-year.

The estimating equation is derived by adding time dummies to equation (2), which eliminates the log market share of used goods on the left-hand side, and controls for substitution between new and used vehicles (Knittel and Stango, 2008):

\[ \ln s_{jt} = \alpha f_{jt} + \tau_t + \phi_{jy} + \nu_{jt}. \]  

(3)

The model-year intercepts, \( \phi_{jy} \), control for characteristics that do not change within the model-year, such as engine characteristics.

Expected fuel costs are:

\[ f_{jt} = \sum_{s=t}^{T+t} \left[ \frac{1}{(1 + r)^s} \frac{P^g_s}{MPG_{jy}} M_s \right]. \]  

(4)

Fuel costs in period \( s \) equal the number of miles driven, \( M_s \), multiplied by the cost of driving one mile, \( P^g_s / MPG_{jy} \), where \( P^g_s \) is the expected price of gasoline in period \( s \) and \( MPG_{jy} \) is fuel efficiency, in miles per gallon (MPG). Total expected fuel costs, \( f_{jt} \), equal discounted expected fuel costs, with a discount rate of \( r \) and a vehicle life of \( T \) periods.

We assume that the price of gasoline follows a random walk, so that the expected price at time \( s > t \) is equal to the price at time \( t \). As a result, the expected cost of driving a specific model
is proportional to the current price of gasoline, divided by the vehicle’s fuel efficiency. We use equation (4) to replace \( j_{jt} \) in equation (3):

\[
\ln s_{jt} = \alpha \frac{P_{t}^g}{MPG_{jy}} + \tau_{t} + \phi_{jy} + \nu_{jt}.
\]  

Equation (5) is the baseline estimating equation. The dependent variable is the log share of sales of model \( j \) in month \( t \). The first independent variable, referred to as dollars-per-mile, is the expected cost of driving the vehicle one mile at the time of purchase; \( P_{t}^g \) is the seasonally adjusted price of gasoline; and \( MPG_{jy} \) is the fuel efficiency of model \( j \) in model-year \( y \). The next section describes the details of the variable construction.

The coefficient of interest is \( \alpha \), which is proportional to the effect of the cost of driving one mile on the log market share. The parameter \( \alpha \) is identified by time-series variation of the price of gasoline and cross-sectional variation of fuel efficiency. In particular, within-year variation of the price of gasoline differentially affects expected driving costs across models. For example, when the expected price of gasoline increases, the fuel costs of a fuel-efficient vehicle increase by less than the costs of a “gas guzzler”.

The model-year intercepts and monthly gasoline price variation are central to the empirical strategy. The model-year intercepts account for the potentially endogenous relationship between the average retail price and vehicle characteristics (Nevo, 2000). They also allow for a different coefficient on the average retail price for each model, thus avoiding the independence of irrelevant alternatives (IIA) assumption that is commonly made in the standard logit model. Note that in equation (5), the assumption that \( \alpha \) is constant across models has different implications than a model in which the coefficient on the retail price is constant across models. Assuming that \( \alpha \) is constant implies that when the price of gasoline changes, the difference in the change in
market shares for any two models is proportional to the difference in the inverse of the models’ fuel efficiencies. For example, an increase in the price of gasoline would cause SUV consumers to substitute from large to small SUVs in the same proportion as would minivan consumers substitute from large to small minivans. This seems to be a more reasonable assumption than the IIA property, but will be relaxed below by allowing $\alpha$ to vary over time and across models.6

The final assumption is that the price of gasoline is exogenous to within model-year determinants of sales. Copeland et al. (2005) and Corrado et al. (2006) have documented that transaction prices (the prices paid by the consumer, as distinct from the manufacturer suggested retail price, MSRP) vary substantially within the model-year. Prices decline dramatically over the course of the model-year, and sales follow a “hump-shaped” pattern, peaking in the early summer. Furthermore, sales and price profiles vary across market segments (e.g., compact cars), and many manufacturers have recently introduced incentives for specific models (Busse et al., 2007). These patterns are the result of complex inventory and vehicle production optimization problems, and which we assume are uncorrelated with dollars-per-mile. Section 6 provides some evidence in support of this assumption.

The motivation for estimating $\alpha$ is that we can then address how much of the changes in market shares in Figure 1 are due to the increase in gasoline prices and to estimate the effect of the gasoline tax on average fuel efficiency. For example, to calculate the effect of the price increase on SUV market shares, we use equation (5) to compute counterfactual market shares in 2007 using the price of gasoline in 2002; the effect of the price increase can be inferred from the differences between actual and counterfactual market shares. Note that the denominator in the

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6 Equation (5) includes several additional functional form assumptions: the vehicle price and driving costs are linear and separable in the utility function, there are no income effects, and the effect of an increase in fuel costs on utility is the same across all consumers. These assumptions are also relaxed below.
market share in equation (5) includes new and used vehicles. We are interested in the effect of gasoline prices on new vehicle market shares, and therefore renormalize market shares to sum to one when calculating the counterfactuals.

Before presenting the results, we argue that \( \alpha \) is the appropriate parameter for answering these questions. As just discussed, we assume that the price of gasoline is exogenous to other determinants of new vehicle demand. However, it is important to note that we allow for the possibility that the price of gasoline affects new vehicle prices, as documented by Langer and Miller (2008). In other words, \( \alpha \) describes the equilibrium relationship between the price of gasoline and new vehicle sales. An increase in the price of gasoline causes a relative outward shift of the demand curves of fuel efficient vehicles and a movement along the corresponding supply curves. The same price increase would cause a relative inward shift of the demand curves for gas guzzlers. Note that both the market share and gas tax questions depend on the relationship between the price of gasoline and equilibrium sales. Estimating \( \alpha \), as opposed to analyzing vehicle prices, is therefore sufficient for answering these questions.

However, these questions are long-run in nature. Because we utilize monthly data to estimate equation (5), the estimate of \( \alpha \) corresponds to the short-run effect of the price of gasoline. Thus, \( \alpha \) depends on how much the demand curves for each vehicle model shift in the short run and on the shape of the corresponding short-run supply curves. If supply curves have the same shapes and demand curves shift by the same amounts in the long run as the short run, the effect of the price of gasoline on new vehicle demand would not vary across time horizons. Of course, this need not be the case. For example, see Bresnahan and Ramey (1993) and Copeland and Hall (2005) on production and pricing decisions in the short and long run. Section 5 assesses the
relationship between the short- and long-run effects of the price of gasoline on new vehicle sales. We find the two to be fairly similar, validating our approach.\(^7,8\)

4 Data

We construct the real price of gasoline with data from the Bureau of Economic Analysis (BEA), using the monthly consumer price index (CPI) and the price of gasoline from 1970-2007.\(^9\) The real price of gasoline, \(P_t^g\), is the price of gasoline divided by the CPI, with the CPI normalized to one in April of 2008. The price of gasoline is seasonally adjusted using X-12 ARIMA, which is the same model used by the Census Bureau.

Model sales are from weekly publications of Wards Automotive Reports (1970-1979) and Ward’s AutoInfoBank (1980-2007). Note that the 1970s sales data do not include light trucks or imports. We match monthly sales data by individual model from 1970 to 2007 to model characteristics data.\(^10\) The characteristics data are available in print in the annual Ward’s Automotive Yearbooks (1970-2007), and include wheelbase, curb weight, engine size

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\(^7\) The price response may vary across firms if an increase in the price of gasoline relaxes the constraint imposed by the CAFE standard (Jacobsen, 2007). Estimating a separate \(\alpha\) for each model-year allows for this possibility, and the results are unaffected (see Table 4).

\(^8\) From the model in Section 2, \(\alpha\) is proportional to miles driven, which is defined as the number of miles driven per year, conditional on vehicle choice. The price of gasoline may affect miles driven, but consumers should account for this effect when making their purchase decisions. Consequently, \(\alpha\) includes the indirect effect of the price of gasoline on sales via miles driven.

\(^9\) More specifically, the gasoline price is constructed from the price of regular unleaded gasoline, the price of regular leaded gasoline, and the nominal price of crude oil. The gasoline price equals the price of unleaded gasoline when the price is available, from 1976-2007. We impute the gasoline price before 1976 using the estimated relationship between the prices of unleaded gasoline, leaded gasoline and oil when the price series overlap.

\(^10\) The match is not straightforward because the two data sets are reported at different levels of aggregation. Vehicle characteristics data are reported at the “trim level” to recognize differences in the MSRP; for example, the data distinguish the 2- and 4-door versions of the Honda Accord sedan. We aggregate the characteristics data to match the model-based sales data, and calculate four statistical moments for the distribution of the vehicle characteristics by car line: minimum, maximum, mean and median. We use the mean value to estimate equation (5), but obtain similar results using other moments.
(displacement) and retail price. Note that the model-year is defined to begin in September for each model.\textsuperscript{11}

The fuel efficiency of each vehicle is obtained from the EPA for model-years 1978-2007.\textsuperscript{12} Fuel efficiency from 1970-1977 is imputed from the estimated relationship between fuel efficiency, wheelbase, weight and engine size from 1978-1979. The imputation is valid under the assumption that vehicle technology did not change in the 1970s, and may introduce some measurement error.

Figure 2a shows the real price of gasoline and the sales-weighted average MPG from 1970-2007, plotting quarterly averages for clarity. Both series vary considerably over time. The price of gasoline increased sharply in late 1973 and again in 1979, coinciding with the major oil shocks, and declined significantly in the mid 1980s. The price was then relatively stable through the 1990s, before increasing from 2002-2007. Average MPG declined gradually in the early 1970s and began increasing in the late 1970s and 1980s, as the CAFE standard, which took effect in 1978, increased.\textsuperscript{13} Fuel efficiency declined steadily in the 1990s, remaining roughly constant thereafter. Much of the decrease was due to compositional changes, particularly the increase in sales of SUVs, which were subject to the lower CAFE standard for light trucks.

Figure 2b shows the log price of gasoline and log average fuel efficiency after taking first differences and removing year and quarter fixed effects. Within-year increases in the price of gasoline are associated with increases in the average fuel efficiency of new vehicles, particularly

\textsuperscript{11} Accounts in the trade press suggest that the first month of the model-year varies somewhat across models in the data, particularly in recent years. The data do not allow us to observe the first month directly, but the specification that includes vehicle class-month interactions partially addresses potential bias that would arise if the first month of the model-year is not random.

\textsuperscript{12} The 1978 Energy Tax Act required that mileage ratings be reported. Note that a previous version of the paper used a fuel efficiency variable constructed from Wards, rather than the EPA, but the latter appears to be more accurate; overall, the differences across data sources are quite small.

\textsuperscript{13} There is a sharp decline in MPG in January of 1980 because the 1970s data do not include light trucks or imports. This change in coverage should not bias the estimation results because of the model-year intercepts.
towards the end of the sample. The model-level estimates of equation (5), discussed in the next section, reflect this pattern.

Table 1 reports additional summary statistics. The first two rows of Panel A show the mean and standard deviation of the monthly observations of the price of gasoline and model-by-month observations of MPG, by decade. The real price of gasoline was lower and less volatile in the 1990s than in other time periods. The MPG distribution is fairly stable, although the share of models with high fuel efficiency has increased gradually over time, as shown in Figure 3.

The last two rows in Panel A of Table 1 show dollars-per-mile and the log of model sales. Panel B shows the standard deviations of the dollars-per-mile and sales variables after taking first differences, which is the transformation used to estimate equation (5) below. Even though this transformation removes much of the variation, considerable variation remains for both variables in all four decades.

5 MAIN RESULTS AND IMPLICATIONS FOR MARKET SHARES AND A GASOLINE TAX

5.1 EFFECT OF THE PRICE OF GASOLINE ON VEHICLE DEMAND OVER TIME

Table 2 reports the estimate of $\alpha$ in equation (5). The dependent variable is the log share of sales by model and month. The main independent variable is dollars-per-mile, defined as the real price of gasoline divided by the fuel efficiency of the model, in MPG. The specification includes time dummies and model-year interactions and the parameters are estimated by Ordinary Least Squares (OLS). The first row of column 1 of Table 2 reports the estimate of $\alpha$, the coefficient on dollars-per-mile. Standard deviations are in parentheses, clustered by model. The estimate is -12.64 with standard error 2.52, which is significant at the one percent level.
The autocorrelation function of the residuals from this specification indicates significant serial correlation, however. For that reason, column 2 reports a first differenced specification of equation (5). The point estimate of $\alpha$ is similar, -10.45, and is again significant at the one percent level. All remaining specifications reported in the paper transform the variables by taking first differences. Serial correlation is further addressed below when lags of the dependent variable are added.

There are several ways to interpret the magnitude in column 2. First, consider the 2003 Acura CL (23 MPG) and the 2003 Volkswagen Jetta (32 MPG). The estimate implies that a one dollar increase in the price of gasoline would reduce sales of the Acura by about 12 percent compared to sales of the Jetta. The second interpretation of the coefficient is in terms of the elasticity of the average MPG of new vehicles with respect to the price of gasoline. The estimated elasticity is about 0.12, which falls between the estimates of Austin and Dinan (2005) and Small and Van Dender (2007).\(^{14}\)

While our primary interest is in explaining the trends shown in Figure 1 and estimating the effect of the gasoline tax on average fuel efficiency, it is first necessary to address whether $\alpha$ has changed over time. If it has, we would restrict the sample to the most recent time period to answer these questions; alternatively, a longer sample period could increase efficiency. Column 3 separates the sample into three periods: 1970-1985, which includes the two oil shocks, the imposition of the CAFE standards and the gasoline price decrease in the mid 1980s; 1986-2001, during which the price of gasoline was nearly constant; and 2002-2007, which includes several

\(^{14}\) Small and Van Dender (2007) and Hughes et al. (2008) provide evidence that the short and long run own price elasticity of gasoline consumption has decreased in magnitude over the past 30 years. The elasticity can be decomposed into three effects: the elasticity of miles travelled with respect to the gasoline price; the price elasticity of the size of the vehicle stock; and the price elasticity of the average fuel efficiency of the stock. The results reported in Small and Van Dender (2007) indicate that the decrease in the own price elasticity of consumption is due to a decrease in the gasoline price elasticity of miles traveled, and not due to a decrease in the elasticity of average fuel efficiency of the vehicle stock. The results in this paper are thus not inconsistent with the previous study.
sharp increases in the price of gasoline, comparable in magnitude to the 1970s shocks. Note that the table reports the interactions of dollars-per-mile with the time period dummies, so each coefficient can be interpreted as the response during the corresponding period (i.e., there is no omitted time period). The results suggest that consumer demand responds to the price of gasoline when the price is high or increasing, in the 1970s and early 1980s and in the most recent period. The price of gasoline had a negligible effect on sales when prices were stable in the middle period, which is consistent with casual observation of the new vehicles market. The results are also similar if dollars-per-mile is interacted with polynomials of a time trend (not reported). In most of the remaining analysis, because we are interested in understanding recent changes in market shares and the effect of a future increase in the gasoline tax, we restrict the sample to the most recent time period (2002-2007). Also note that we return to the possibility of asymmetric demand responses at the end of Section 5.2.

5.2 SHORT AND LONG RUN EFFECTS OF GASOLINE PRICES

Table 2 shows that monthly gasoline prices have a statistically and economically significant effect on monthly new vehicle sales. Yet the two issues discussed in the introduction, the changes in market shares between 2002 and 2007 as well as the potential impact of an increase in the gasoline tax on overall fuel economy, are related to longer-term effects of gasoline prices on vehicle sales. This section shows that the short and long run effects of a change in the price of gasoline are probably fairly similar in magnitude.

15 We have also estimated specifications that separate the first time period into 1970-1978 and 1978-1985. There is a much stronger demand response in the latter sub-period, which we believe is at least partially due to measurement error (recall that MPG is imputed before 1978). In support of this argument, the estimates are smaller for the latter periods if we use the imputed MPG rather than the actual MPG. Because our primary interest is in characterizing consumer demand in the most recent time period, we do not pursue this question further.
In principle, the short run effect could be larger or smaller in magnitude than the long run. On the one hand, if vehicle prices are sticky, the short run quantity effect would be larger. On the other hand, short-run production constraints or lags in consumers’ demand response could cause the long run quantity effect to be larger. For example, if prices increase in the short run to clear the market because of production constraints, \( \alpha \) would understate the long run effect of the price of gasoline on demand.

Column 1 of Table 3 reports the same specification as column 3 of Table 2, except that the sample is restricted to the model-years 2002-2007. The remainder of the paper refers to this specification as the baseline. Columns 2-8 in Table 3 report specifications that investigate the relationship between the short and long run effects. Columns 2-4 add three lags of dollars-per-mile, three lags of the dependent variable, and 3 lags of both variables. Column 2 shows that the one-month lag of dollars-per-mile has a moderate, although not statistically significant, effect on sales, and additional lags are not significant. Column 3 shows that adding lags of the dependent variable does not affect the estimate on dollars-per-mile, although some of the lags are statistically significant. The coefficient on the lagged variable implies that a one percent decrease in last month’s sales is associated with a 16 percent increase in current sales. The negative sign suggests that a gasoline price shock causes automobile dealers to draw down inventories, after which vehicle prices respond to partially offset the quantity change. The result is consistent with Langer and Miller (2008), who find that transaction prices respond to current and lagged gasoline prices. These estimates can be used to calculate the within-year effect of the price of gasoline on sales (Small and Van Dender, 2007); based on the estimates for the lag dependent variables, the within-year response is therefore within about 20 percent of the within-month
response. If we consider the within-year effect to constitute the medium run, we conclude that the short and medium run quantity responses are similar in magnitude.

It is also possible that the effect across years is different from the within-month or within-year effects. For example, production constraints could explain such long lags – several automakers have recently announced plans to significantly reduce production of large SUVs over the next five years. We take two approaches to investigate this possibility. First, we construct a variable that proxies for the change in consumer demand over the previous year for the vehicle. The variable is motivated by a model in which firms plan production in August for the following model-year based on the expected price of gasoline. Consider production planning in August of year \( y - 2 \). If the price follows a random walk, the error in the average expected price over the following model-year equals the difference between the average price (September to the following August) and the price in August: \( \bar{p}_{y-1} - p_{Aug, y-2} \). The unexpected decrease in average demand for the vehicle over the model-year is therefore proportional to the ratio of the expectation error to the fuel efficiency of the vehicle: \( \frac{\bar{p}_{y-1} - p_{Aug, y-2}}{MPG_{j, y-1}} \); we define this variable as the lag demand shock. If a firm experiences a decrease in demand for its vehicle over model-year \( y - 1 \), then by the random walk assumption, this decrease should be permanent. Therefore, the firm will reallocate production in model-year \( y \). If the lag demand shock is added to the baseline equation, we expect the coefficient on the variable to be negative. Columns 5-8 show that the main results are not affected much by adding this variable to the specifications in columns 1-4. The coefficient on the lag demand shock is negative in the lag dependent variable specifications, but the magnitude is much smaller than \( \alpha \). There is thus some evidence that firms
re-allocate production across years in response to unexpected changes in the price of gasoline, but the average effect is not quantitatively large.

The second approach to investigating the long run response is to use model dummies instead of model-year interactions in equation (5) and include additional lags of dollars-per-mile. If 36 lags are added to equation (5), the coefficient on current dollars-per-mile is similar to the baseline specification. Although there is some evidence for a lagged effect of up to two years, the magnitudes are relatively small. The results are not reported due to space considerations, but are available upon request.

Table 2 showed that the effect of the price of gasoline on sales has varied over time. This pattern raises the possibility that the relationship between gasoline prices and sales is nonlinear or asymmetric. In the baseline specification in column 1 of Table 3, $\alpha$ is the average effect of dollars-per-mile on log market share. If the effect is greater at high gasoline prices, or differs in magnitude for price increases and decreases, the estimates would lead to incorrect conclusions about the causes of recent changes in market shares or the effect of a gasoline tax. Column 9 adds to the baseline equation the interaction of dollars-per-mile with a dummy variable that is equal to one if the price of gasoline increased between the previous and current month (the time dummies absorb the main effect of the dummy variable). The effect of a price increase on sales would be larger than a decrease if the interaction were negative, but the results indicate that the response is roughly symmetric. We have investigated other specifications that allow for nonlinear or asymmetric responses, such as following Dargay and Gately (1997) by decomposing price changes into three series: the maximum price up until the current month, a non-decreasing series of price increases and a non-increasing series of price cuts.\footnote{More precisely, the price increase series equals the cumulative sum of price decreases since the first month in the sample, where a price increase at time $t$ is the difference between the price at time $t$ and the price at time $t - 1$ if the}
specification distinguishes between rising and falling prices and between short and long-term changes. The results in column 10 do not provide evidence that demand responds asymmetrically or in a non-linear fashion, as the coefficients on the three variables are statistically indistinguishable. Overall, we find the estimate of $\alpha$ reported in column 1 of Table 3 to be close to several different estimations of what we consider to be the long run effect of the price of gasoline on new vehicle demand. While this result is perhaps somewhat surprising, the upshot is that we can use the empirical approach to address the two main questions: the cause of recent changes in market shares and the effect of the gasoline tax on average fuel efficiency.\footnote{We have also estimated specifications that aggregate monthly sales and gasoline prices to quarterly sales and prices. If the long run response were different in magnitude than the monthly response, the quarterly results would likely be different. The estimates are available upon request, and are similar to the baseline.}

5.3 Effect of Gasoline Prices on Market Shares of U.S. Firms and SUVs

Between August of 2002 and August of 2007 the real price of gasoline increased from $1.75 to $2.86 per gallon. During the same time period, the market share of small SUVs increased from 7.5 to 10.5 percent, while the market share of large SUVs decreased from 18.3 to 10.5 percent. At the same time, market shares of U.S. manufacturers, which rely on sales of large vehicles, have declined by 20 percent. It is unclear, though, how much of these changes have been caused by gasoline prices.

Our identification strategy allows us to determine how much of the recent changes in market shares are caused by the increase in the price of gasoline. We assume that within-year changes in gasoline prices are uncorrelated with within-year changes in preferences, i.e., that the model-year intercepts absorb changes in preferences. This assumption is valid if preferences are slow-
moving and within-year preference shocks are uncorrelated with the price of gasoline, which seems likely to be the case; Section 6.2 provides some empirical support for this assumption. We estimate the effect of the price increase on SUV market shares by computing the counterfactual market shares if the price had remained constant at the 2002 level. We then compare the actual and counterfactual SUV market shares to determine the effect of gasoline prices on large SUV demand, and similarly for small SUVs. The estimate of $\alpha$ in column 1 of Table 3 implies that the increase of $1.11$ per gallon caused about 25 percent of the decrease in the market share of large SUVs and the increase in the market share of small SUVs.

Turning to U.S. firms, about 40 percent of the decrease in U.S. market share has been caused by the recent increase in the price of gasoline. We conclude that the price of gasoline has a substantial effect on the new vehicles market, although perhaps smaller than some recent analysts have suggested.

5.4 Effect of a Gasoline Tax on Average Fuel Efficiency

We now relate our model to the economic policy issue of reducing gasoline consumption. The policy debate has focused on raising the CAFE standard, a command-and-control type regulation that applies to new vehicles sold. Many economists, however, have argued that raising the gasoline tax instead would be more efficient. The welfare comparison of the two policies depends partly on the sensitivity of new vehicle demand to the price of gasoline (Austin and Dinan, 2005). Equation (5) estimates precisely this effect.

---

18 The negative estimate of $\alpha$ implies that all market shares decrease when the price of gasoline increases. It is therefore necessary to renormalize the market shares to sum to one when comparing the actual and counterfactual SUV market shares. A similar renormalization is made in the following discussion of the gasoline tax.

19 These results do not appear to be sensitive to the functional form assumptions in equation (5), as the results are similar if a separate $\alpha$ is estimated for each model-year.
We use the estimate of $\alpha$ to calculate the change in average fuel efficiency of new vehicles due to a one dollar increase in the price of gasoline. The calculation is based on the predicted market shares of models for which sales are positive at the end of the sample period. The first column of Table 4 reports the difference between the predicted and actual sales-weighted MPG for the baseline specification in column 1 of Table 3. The standard error is reported in parentheses, calculated using the delta method.\textsuperscript{20} The estimate of $\alpha$ implies that a one dollar price increase would raise average fuel efficiency by 1.08 MPG, which is significant at the one percent level. As noted above, the elasticity of average fuel efficiency with respect to the price of gasoline is therefore about 0.12, which is roughly one-half of that reported by Austin and Dinan.

6 ROBUSTNESS CHECKS

6.1 ALTERNATIVE ESTIMATION MODELS

As noted above, equation (5) includes several functional form assumptions. Columns 2-4 of Table 4 report the estimated effect of a one dollar price increase on average fuel efficiency using a number of different estimation models that relax the main assumptions in equation (5).

Equation (5) was derived by assuming that dollars-per-mile has the same effect on every consumer’s utility. By comparison, random coefficients logit models, such as BLP, allow for a separate $\alpha_i$ for each consumer. In that case, the effect of dollars-per-mile on the market share of a particular model is the average $\alpha_i$, weighted by the probability individuals purchase the specific model. Consequently, there is a separate $\alpha_j$ for each model. Researchers commonly assume that $\alpha_j$ is normally distributed and estimate the mean and standard deviation of the distribution. Our

\textsuperscript{20} We assume that actual market shares are measured without error, and that the only uncertainty arises from sampling variation over $\alpha$. The standard errors are computed by taking the first order approximation of the mean market share, which is a nonlinear function of $\alpha$. 


23
data include sufficient observations to simply estimate a separate $\alpha_{jt}$ for each model-year in equation (5) using observations from 2002-2007. Figure 4 plots a histogram of the estimated coefficients, which indicates that there is some heterogeneity, but that most coefficients fall within a fairly narrow interval, between -5 and -30. Column 2 in Table 4 uses the estimated coefficients from this specification to calculate that a one dollar price increase raises average fuel efficiency by 1.20 MPG, which is quite similar to the baseline in column 1 and is significant at the 5 percent level.

We next estimate a model based loosely on the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer (1980), which gives an arbitrary first order approximation to any set of demand equations. In the standard AIDS model, the dependent variable is the model’s share of revenue in total market revenue and the independent variables are the prices of all models. But in the current setting, the total cost of purchasing and using a model is the sum of the vehicle’s price and operating costs. As in equation (5), including model-year intercepts absorbs the vehicle’s price and maintenance costs, leaving the following equation:

$$w_{jt} = \gamma_j \ln P_{jt} + \lambda_j \ln \hat{Q}_t + \phi_{jt} + \nu_{jt},$$

where the dependent variable is the revenue share of the model calculated using annual retail prices. The first independent variable is the log price of gasoline; the second independent variable is aggregate sales, using the approximation in Deaton and Muellbauer; $\phi_{jt}$ is a model-year fixed effect and $\nu_{jt}$ is an error term. The coefficient $\gamma_j$ is the effect of a one percent increase in the price of gasoline on the revenue share of model $j$; the specification allows the price of gasoline to affect the revenue share of each model differently. The second independent variable controls for the relationship between total sales and the revenue share of each model, and the coefficient can also vary freely across models. We estimate equation (6) by taking first
differences of all variables using observations from 2002-2007. Figure 5a shows a histogram of the \( \gamma_j \) coefficients and Figure 5b plots the coefficients against the MPG of the corresponding model. Figure 5b shows a strong positive correlation, indicated by the fitted line, which suggests that an increase in the price of gasoline causes the relative revenue shares of fuel efficient vehicles to increase. Column 3 of Table 5 confirms this result, using the predicted change in market shares to calculate the change in average MPG caused by a one dollar gasoline price increase. The estimate is 0.79 MPG, which is somewhat smaller than the other estimates. The estimate is not statistically significant, which is not surprising given the large number of estimated coefficients.

Finally, instead of making assumptions on consumer preferences and the distributions of unobserved parameters, we can estimate an aggregate regression that characterizes the effect of the price of gasoline on average fuel efficiency:

\[
\ln MPG_t = \delta \ln P_t^g + \mu_m + \tau_j + \omega_t, \tag{7}
\]

The dependent variable is the log of the monthly sales-weighted average MPG of new vehicles and the first independent variable is the log monthly price of gasoline. The regression includes month and year dummies and the coefficient \( \delta \) is the elasticity of average MPG to the price of gasoline. The advantage of this specification is that \( \delta \) is simple to interpret, as a linear approximation to the effect of the gasoline price on average MPG. On the other hand, the equation cannot be derived from a model of consumer behavior and the results cannot be used to answer questions pertaining to the effect of gasoline prices on market shares of U.S. manufacturers and SUVs. Nevertheless, equation (7) is estimated for comparison using observations from 2002-2007. The estimate of \( \delta \) is 0.063 with standard error 0.015, which is significant at the one percent level, implying that a one percent increase in the price of gasoline...
raises average fuel efficiency by 0.06 percent. Column 4 of Table 4 reports that a one dollar increase in the price of gasoline raises average fuel efficiency by 0.50 MPG, which is smaller than the other estimates, but which has similar implications for the effect of the gasoline tax.

6.2 Additional Specifications

The preceding sections have documented a strong relationship between the price of gasoline and the demand for fuel efficient vehicles. Next we utilize alternative vehicle characteristics. Models with greater fuel efficiency are generally smaller and lighter, so an increase in the price of gasoline should also lead to larger market shares of small and light models. Columns 1-4 of Table 5 show that this is the case, which provides further support for the main results.

We consider four characteristics: weight, length, engine size and number of cylinders. The fact that driving costs should affect consumer decisions similarly to model prices motivated the use of dollars-per-mile when analyzing fuel efficiency, but there are no corresponding theoretical arguments to guide the functional forms for the other characteristics. Instead, we use a semi-parametric approach and separate models into quartiles for weight, length and engine size, as well as three categories for engines with 4, 6 or 8 cylinders. Columns 1-4 report estimates of equation (5), replacing dollars-per-mile with the interactions of the log price of gasoline with the quartile or cylinder indicator variables. Models that are light, short or have small engines constitute the first and omitted categories in the four columns. The results consistently show that an increase in the price of gasoline causes a decrease in the market shares of large, heavy models that have large engines.

In the baseline specification we assume that dollars-per-mile is exogenous to other time-varying determinants of sales, such as transaction prices. As noted earlier, the price of gasoline
could affect transaction prices, which was investigated using a reduced form approach in Table 3. However, the price of gasoline may be correlated with transaction prices or other unobserved variables and therefore would not be truly exogenous. We next investigate bias arising from two potential omitted variables, and conclude that the empirical strategy appears to be robust.

Previous work, e.g., Copeland et al. (2005) has found that transaction prices vary differentially within the model-year for different market segments and sizes. Column 5 controls for this pattern by including the interaction of a set of month dummy variables with a set of indicators for the quintile of the model’s fuel efficiency. This specification allows models in each fuel efficiency quintile to have different transaction price and sales profiles. The estimate is quite similar to the baseline and suggests that these profiles are uncorrelated with dollars-per-mile.21

Another possible source of bias is that the price of gasoline may be correlated with the composition of consumers that purchase vehicles. Sufficiently detailed and high frequency data are not available to investigate this possibility directly, but we can address this concern by adding calendar month-vehicle market segment interactions to the baseline specification. The specification controls for seasonal patterns of demand. For example, if the price of gasoline happens to increase during winter months in the sample and SUV consumers are less likely to purchase vehicles in the winter, the estimate of $\alpha$ could reflect a spurious correlation. Column 6 controls for seasonal effects, and shows that the estimate of $\alpha$ remains robust. The specification provides support for the assumption that the price of gasoline is uncorrelated with consumer preferences.

7 CONCLUSION

21 Manufacturer incentives have become increasingly common in the past decade (Busse et al., 2007), particularly among U.S. manufacturers. These incentives do not appear to bias the results, as the estimates are similar if we omit U.S. firms.
This paper estimates the effect of the price of gasoline on the demand for fuel efficient vehicles. The empirical strategy combines time series variation of the price of gasoline with cross sectional variation of fuel efficiency, exploiting the fact that the effect of a gas price change on fuel costs is inversely proportional to fuel efficiency. We control for unobserved characteristics that vary by model-year by using monthly gasoline price and sales data, combined with model-year fixed effects. The price of gasoline has a significant effect on the demand for fuel efficient vehicles. Based on specifications including lags and allowing for asymmetries, we find that the short run quantity response to a price shock is similar to the long run response. The estimates imply that the increase in the price of gasoline from 2002-2007 explains much of the change in the market shares of SUVs and of U.S. automakers. Turning to the policy question of using the gasoline tax to reduce fuel consumption, an increase in the Federal gasoline tax that raises the price of gasoline by one dollar would raise the average fuel efficiency of new vehicles by about 0.5-1 MPG.

This study, as well as previous empirical work, considers the effect of the price of gasoline on new vehicle demand, using a static approach in which the set of models is exogenous. Gasoline prices and regulations such as CAFE may affect the characteristics of vehicles in the market, including fuel efficiency. Further work should consider a dynamic setting, in which vehicle characteristics are endogenous.

8 REFERENCES

8. Copeland, Adam, Wendy Dunn and George Hall (2005), “Prices, Production and Inventories Over the Automotive Model Year.”
27. Ward’s Automotive Yearbook, 1980-2003, Ward’s Communications
28. Ward’s AutoInfoBank, Ward’s Automotive Group
Figure 1: Change in Market Shares of U.S. Firms and SUVs and Gasoline Price, 2002-2007

Notes: U.S. firms include Chrysler, Ford and GM. Small sport utility vehicles (SUVs) include crossover utility vehicles, and large SUVs include all other SUVs. The figure plots the change in the log quarterly average of the real gasoline price and market shares of U.S. firms and SUVs, relative to the first quarter of 2002. The real price of gasoline is computed from the BLS and market shares are computed from Wards Auto. See Section 4 for details on data construction.
Notes: Average miles per gallon (MPG) is the sales-weighted average MPG by year and quarter. Figure 2a plots average MPG from 1970-2007, and Figure 2b plots the residual of the log of average MPG, after taking first differences and removing annual means and quarter fixed effects. The real price of gasoline is the price of unleaded gasoline divided by the consumer price index, using the national average gasoline price and the consumer price index (CPI) from the Bureau of Labor Statistics. The CPI is normalized to one for April, 2008. Figure 2a plots the real price of gasoline, in 2008 dollars, and Figure 2b plots the residual of the log real price, constructed similarly to average MPG (see text for details).
Notes: The figure plots a histogram of the fuel efficiency, in MPG, of vehicles in the Wards database in the indicated years. The vertical axis is the share of models with positive sales for the corresponding year. The horizontal axis labels the maximum fuel efficiency of vehicles in the bin.
Figure 4: Histogram of Coefficients on Dollars-per-mile by Model

Notes: Figure 4 reports a histogram of the estimated coefficients from equation (5) with a separate coefficient on dollars-per-mile for each model-year. The sample, dependent variable and other independent variables are the same as in column 1 of Table 3. The histogram shows the number of models for which the estimated coefficient falls within the indicated bin.
Notes: Figure 5a reports a histogram of the estimated coefficients from estimating equation (6) separately by model. The dependent variable is the share of revenue in total revenue by month and year for the particular model. The independent variables are a set of year dummies, the log real price of gasoline and the log of aggregate sales. The histogram shows the number of models for which the estimated coefficient falls within the indicated bin. Figure 5b plots the coefficients against the MPG of the model. The solid line indicates the fitted values of an OLS regression of the coefficient on MPG.
Table 1:

Summary Statistics

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<td><strong>Panel A: Sample Means and Standard Deviations</strong></td>
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<td>2.41</td>
<td>1.65</td>
<td>2.23</td>
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<td></td>
<td>(0.18)</td>
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<td>(0.16)</td>
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<td>MPG</td>
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<td>(2.88)</td>
<td>(5.04)</td>
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<td>(1.07)</td>
<td>(1.57)</td>
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<td><strong>Panel B: Standard Deviations After First Differencing</strong></td>
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Notes: Cells in Panel A report means with standard deviations in parentheses. The first row of Panel A reports the monthly real gasoline price, computed as in Figure 2. The second row reports the average MPG of all models sold in the indicated decade. The third row reports dollars-per-mile, defined as the ratio of the price of gasoline to MPG. The fourth row reports log monthly sales by model. Panel B reports the standard deviation of the indicated variables after first differencing by model-year.
### Table 2: Effect of the Price of Gasoline on New Vehicle Sales, 1970-2007

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<td>-10.10</td>
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<td>(3.48)</td>
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<tr>
<td>Dollars-Per-Mile x 1986-2001</td>
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Notes: Standard errors are in parentheses, clustered by model. The table reports the results of estimating equation (5) by Ordinary Least Squares (OLS). The dependent variable is the log share of sales by model and month. All variables are in first differences in columns 2 and 3. All specifications include month dummies and column 1 includes model-year interactions. Columns 1 and 2 report the estimated coefficient on dollars-per-mile, which is constructed as in Table 1. Column 3 reports the interaction of dollars-per-mile with a set of dummy variables, which are equal to one in the indicated years.
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<td>-0.02</td>
<td>-0.02</td>
<td></td>
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<tr>
<td><strong>Dependent Variable</strong></td>
<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.03)</td>
<td>(0.03)</td>
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<tr>
<td><strong>Three Month Lag</strong></td>
<td>0.04</td>
<td>0.04</td>
<td>0.03</td>
<td>0.04</td>
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<td></td>
</tr>
<tr>
<td><strong>Dependent Variable</strong></td>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.02)</td>
<td>(0.02)</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td><strong>Lag Demand Shock</strong></td>
<td>1.04</td>
<td>1.10</td>
<td>-3.17</td>
<td>-2.77</td>
<td></td>
<td></td>
<td></td>
<td></td>
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<td></td>
</tr>
<tr>
<td></td>
<td>(1.64)</td>
<td>(1.63)</td>
<td>(2.30)</td>
<td>(2.32)</td>
<td></td>
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</tr>
<tr>
<td><strong>Price Increase x</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.35</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td><strong>Dollars-Per-Mile</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(0.25)</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td><strong>Max Price-Per-Mile</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>-9.07</td>
<td>(8.22)</td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td><strong>Price Cut-Per-Mile</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>-9.70</td>
<td>(2.54)</td>
<td></td>
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<tr>
<td><strong>Price Increase-Per-Mile</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>-11.44</td>
<td>(2.96)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>N</strong></td>
<td>15,810</td>
<td>11,493</td>
<td>11,214</td>
<td>11,214</td>
<td>15,189</td>
<td>15,189</td>
<td>10,902</td>
<td>10,902</td>
<td>15,810</td>
<td>68,662</td>
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<tr>
<td><strong>R²</strong></td>
<td>0.02</td>
<td>0.02</td>
<td>0.04</td>
<td>0.04</td>
<td>0.02</td>
<td>0.02</td>
<td>0.04</td>
<td>0.04</td>
<td>0.02</td>
<td>0.02</td>
</tr>
</tbody>
</table>
Notes to Table 3: Standard errors are in parentheses, clustered by model. The dependent variable is the log share of sales, constructed as in Table 2. Column 1 reports the same specification as column 2 of Table 2, restricting the sample to September, 2002-August, 2007. Columns 2-4 report the same specification as column 1, adding the 1-3 month lags of dollars-per-mile or the 1-3 month lags of the dependent variable. Columns 5-8 repeat the specifications from columns 1-4, adding the lag demand shock variable, defined in the text. Price increase is a dummy variable, equal to one if the current price is greater than the price in the previous month. Column 9 adds to the specification in column 1 the interaction of the price increase with dollars-per-mile. Column 10 uses the full sample, 1970-2007, and replaces dollars-per-mile with three variables defined in Dargay and Gately (1997). Max price is the maximum gasoline price between the beginning of the sample and the current month. Price cut is a non-increasing series, equal to the cumulative sum of price decreases since the beginning of the sample. Price increase is a non-decreasing series, equal to the cumulative sum of price increases since the beginning of the sample. Column 10 includes the ratio of max price, price cut and price increase to the model's fuel efficiency. See text for details.
Table 4:

<table>
<thead>
<tr>
<th>Specification</th>
<th>Column (1) of Table 3</th>
<th>Column (1), with Separate Coefficient by Model-Year</th>
<th>Equation (6)</th>
<th>Equation (7)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Effect of One Dollar Price Increase on MPG</td>
<td>1.08 (0.19)</td>
<td>1.20 (0.59)</td>
<td>0.79 (0.56)</td>
<td>0.51 (0.12)</td>
</tr>
</tbody>
</table>

Notes: Each column reports the effect of a one dollar increase in the price of gasoline on average MPG. The effects are calculated from the indicated specifications, which use observations from 2002-2007. Column 1 uses the same specification as column 1 of Table 3. Column 2 uses the same specification as column 1, except that a separate coefficient on the dollars-per-mile variable is estimated for each model-year. Column 3 reports the results of estimating equation (6). Column 4 reports the results of estimating equation (7). In columns 1-3, the calculation uses the predicted market shares of models sold in August, 2007, with and without the price increase. The standard error is in parentheses, calculated using the delta method. The effect of the price increase in column 4 is the change in average miles per gallon if the price increases by one dollar, relative to the price in August, 2007.
## Table 5: Other Model Characteristics and Controlling for Model-Sales Profiles

<table>
<thead>
<tr>
<th>Weight Quartile 2 x Log Price</th>
<th>(1)</th>
<th>0.44</th>
<th>(2)</th>
<th>(0.18)</th>
<th>(3)</th>
<th>(0.18)</th>
<th>(4)</th>
<th>(0.18)</th>
<th>(5)</th>
<th>(0.16)</th>
<th>(6)</th>
<th>(0.16)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Weight Quartile 3 x Log Price</td>
<td>-0.97</td>
<td>0.18</td>
<td>-0.93</td>
<td>0.18</td>
<td>-0.97</td>
<td>0.18</td>
<td>-0.93</td>
<td>0.18</td>
<td>-0.97</td>
<td>0.18</td>
<td>-0.93</td>
<td>0.18</td>
</tr>
<tr>
<td>Length Quartile 2 x Log Price</td>
<td>-0.76</td>
<td>0.18</td>
<td>-0.66</td>
<td>0.17</td>
<td>-0.76</td>
<td>0.18</td>
<td>-0.66</td>
<td>0.17</td>
<td>-0.76</td>
<td>0.18</td>
<td>-0.66</td>
<td>0.17</td>
</tr>
<tr>
<td>Length Quartile 3 x Log Price</td>
<td>-1.07</td>
<td>0.17</td>
<td>-1.07</td>
<td>0.17</td>
<td>-1.07</td>
<td>0.17</td>
<td>-1.07</td>
<td>0.17</td>
<td>-1.07</td>
<td>0.17</td>
<td>-1.07</td>
<td>0.17</td>
</tr>
<tr>
<td>Engine Size Quartile 2 x Log Price</td>
<td>-0.73</td>
<td>0.18</td>
<td>-0.88</td>
<td>0.17</td>
<td>-0.73</td>
<td>0.18</td>
<td>-0.88</td>
<td>0.17</td>
<td>-0.73</td>
<td>0.18</td>
<td>-0.88</td>
<td>0.17</td>
</tr>
<tr>
<td>Engine Size Quartile 3 x Log Price</td>
<td>-1.05</td>
<td>0.17</td>
<td>-1.05</td>
<td>0.17</td>
<td>-1.05</td>
<td>0.17</td>
<td>-1.05</td>
<td>0.17</td>
<td>-1.05</td>
<td>0.17</td>
<td>-1.05</td>
<td>0.17</td>
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<tr>
<td>Engine Size Quartile 4 x Log Price</td>
<td>-0.75</td>
<td>0.18</td>
<td>-1.08</td>
<td>0.17</td>
<td>-0.75</td>
<td>0.18</td>
<td>-1.08</td>
<td>0.17</td>
<td>-0.75</td>
<td>0.18</td>
<td>-1.08</td>
<td>0.17</td>
</tr>
<tr>
<td>6 Cylinder Engine x Log Price</td>
<td>-14.71</td>
<td>2.91</td>
<td>-14.71</td>
<td>2.91</td>
<td>-14.71</td>
<td>2.91</td>
<td>-14.71</td>
<td>2.91</td>
<td>-14.71</td>
<td>2.91</td>
<td>-14.71</td>
<td>2.91</td>
</tr>
<tr>
<td>8 Cylinder Engine x Log Price</td>
<td>-13.51</td>
<td>2.77</td>
<td>-13.51</td>
<td>2.77</td>
<td>-13.51</td>
<td>2.77</td>
<td>-13.51</td>
<td>2.77</td>
<td>-13.51</td>
<td>2.77</td>
<td>-13.51</td>
<td>2.77</td>
</tr>
</tbody>
</table>

**Notes:** Standard errors are in parentheses, clustered by model. The dependent variable is log sales by model and month. The sample is the same as in column 1 of Table 3. For each month and year, models are assigned quartiles based on their curb weight, length (wheelbase) and engine size (liters). The quartile dummies are interacted with the log price of gasoline. Columns 1-3 report the interaction coefficients, where the first quartile is the omitted category. In column 4 models are separated into three categories, depending on the number of cylinders of the engine. The table reports the coefficients on the 6-cylinder and 8-cylinder interactions, where the omitted category has 4 cylinders. Columns 5 and 6 report the same specification as column 1 of Table 3. Column 5 includes calendar month-MPG quintile interactions and column 6 includes calendar month-vehicle class interactions.