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Abstract

Fatal motor vehicle crashes per capita remained relatively stable over the 1990s, in spite of new traffic safety laws and vehicle innovations. One explanation for this stability is that the price of gasoline declined, which resulted in more vehicle miles traveled and potentially more fatalities. By using 1983–2000 monthly gasoline price and fatality panel data and fixed effects specifications, this study exploits within-state variation over time in gasoline prices and found that a 10-cent decrease in gasoline prices increased motor vehicle fatalities by 2.3 percent over a 2-year period. The effect on higher-risk younger adults is more than twice as large. The secular decline in real gasoline prices over the 1990s may partially explain why motor vehicle fatalities per capita have not decreased even with the adoption of more stringent state policies. © 2004 by the Association for Public Policy Analysis and Management.

INTRODUCTION

The Centers for Disease Control and Prevention (CDC, 1999) has hailed the reduction in motor vehicle fatalities as one of the 10 great public health achievements of the 20th century. This success was attributed to safer vehicles, improved roads, greater use of seat-belts, and concerted efforts to reduce drunk driving. Yet, the per capita motor vehicle fatality rate has been relatively stable since the early 1990s, even in the face of several new policy initiatives directed at lowering traffic-related fatalities. One explanation for this stability is that other conditions of driving have changed over this period and effectively lowered the price of driving, which has resulted in more vehicle miles traveled and potentially more fatalities. This study evaluates the role of gasoline prices in the interpretation of aggregate traffic fatality trends.

Motor vehicle fatality rates declined significantly over the last quarter century (see Figure 1). Over the period 1976 through 2000, fatal crashes declined 31.4 percent from 20.88 to 14.87 per 100,000 individuals. On a per vehicle mile traveled (VMT) basis, fatalities declined 52.9 percent from 3.26 to 1.53 per 100 million miles traveled. The greater percentage decline in fatalities per VMT relative to fatalities per capita indicates that, on average, individuals are driving more miles over time. This difference is particularly apparent during the 1990s. Over the period 1992 through 1998, fatalities per capita declined only 0.2 percent, while fatalities per VMT declined 9.7 percent.

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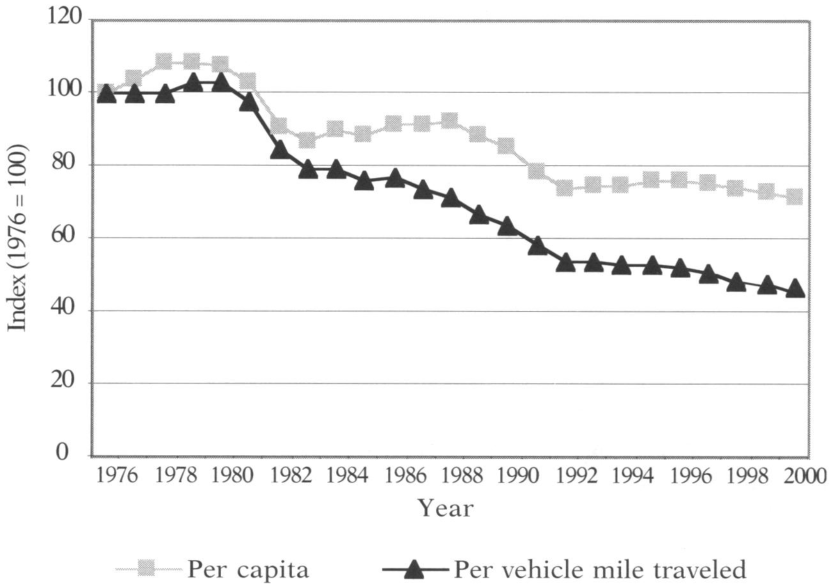


Figure 1. Motor vehicle fatality rates (1976–2000).

This stability in fatalities per capita over the 1990s has been troubling to traffic safety officials (GAO, 2003), especially in light of a number of new policy measures and safety innovations that have been introduced over this period. States have pursued stricter drinking-and-driving laws by the adoption of lower illegal per se blood alcohol concentration (BAC) limits, the revocation of drivers' licenses prior to any court action (i.e., administrative license revocation), and laws that make it illegal for individuals under age 21 to have a positive BAC (i.e., "zero tolerance" laws). Over the period 1992 through 1998, 11 states adopted stricter BAC limits, 12 states adopted administrative license revocation laws, and 31 states adopted zero tolerance laws. In addition, nine states adopted secondary seat-belt laws over this period, under which motorists can be cited for failure to wear a seat-belt if they are pulled over for some other violation. Outside direct policy interventions, vehicle and roadway improvements continued over the 1990s with the introduction of vehicle airbags representing the most prominent safety development. In 1988, 2 percent of new cars had air bags, but by the 1996 model year, 90 percent of all new cars were equipped with this feature (Mannering and Winston, 1995).

One interpretation of the stable per capita fatality rate over the 1990s is that these policies and innovations did not improve overall safety, despite strong evidence that drinking-and-driving laws (Dee, 2001), seat-belt laws (Evans and Graham, 1991), and airbags (IIHS, 1997) have been effective at lowering fatalities. An alternative interpretation, which we explore in this paper, is that other factors have attenuated the potential gains associated with recent state laws and vehicle and road improvements. Two such factors are apparent. A first is the repeal of the national maximum speed limit law (NMSL) in 1995, after which 44 states raised their speed limit. The overall statewide implications of these laws have been debated (Lave and Elias, 1997), but there is evidence that fatalities increase on rural interstates with the adoption of higher speed limits (Greenstone, 2002). Although higher speed limits may have offset some of the recent safety gains, we argue that a primary factor in explaining the stability in fatalities per capita over the 1990s was the low price of gasoline over this period (see Figure 2).

The price of gasoline is potentially related to motor vehicle fatalities through a series of links.[1] The first link involves the relationship between the price of gasoline and the quantity of gasoline consumed. Estimates of the short-run elasticity of demand are in the range of –0.1 to –0.5 (Dahl, 1986, 1995; Dahl and Sterner, 1991; Goodwin, 1992). Long-run elasticity estimates are in the range of –0.6 to –1.2 (Baltagi and Griffin, 1984; Dahl and Sterner, 1991). The high end of these estimates suggests that a 10 percent reduction in the price of gasoline leads to a 12 percent long-run increase in gasoline consumption. The long-run response stems not only from decisions to make an extra trip to the beach, for example, but also from decisions to change long-term contracts and make different capital choices. Shippers may use trains less and long-haul trucks more in the face of sustained lower gasoline prices. People may move to homes more distant from their work or change jobs. Importantly, they may also buy larger, less fuel-efficient cars.

The second link is between gasoline consumption and the number of VMT. In the short run, it is straightforward that an increase in gasoline consumption will translate into more VMT. However, this increase may not be proportional over the long run as individuals substitute toward vehicles with different fuel efficiency levels. In the face of lower gasoline prices, an individual may switch from a mid-sized car to a less fuel-efficient sports utility vehicle (SUV). In this instance, the individual

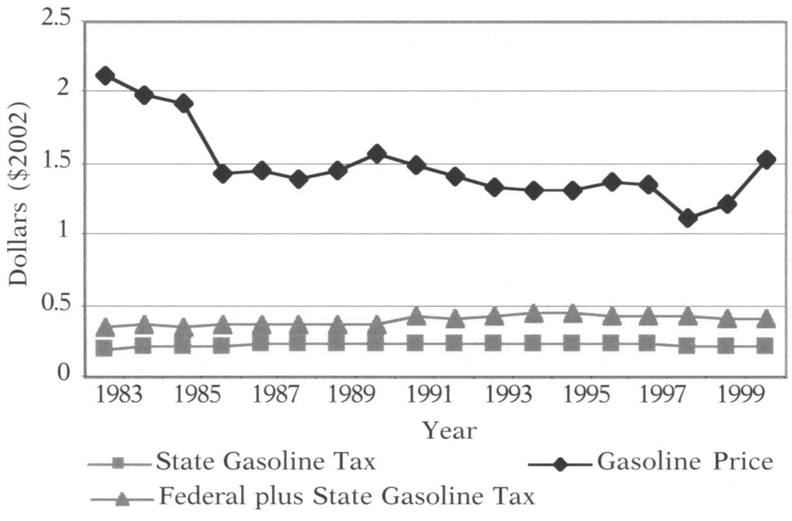


Figure 2. Average gasoline price and taxes per gallon of regular unleaded (1983–2000).

[1] For a detailed discussion of the structural relationships between gasoline prices and motor vehicle safety, see Gaudry and Lasarre (2000).

would consume more gasoline for an equivalent number of VMT. Thus, over the long run, there may be an attenuation of the relationship between gasoline consumption and VMT due to the substitution toward vehicles of varying size.

The final link is between the number of VMT and fatalities. When people drive more miles, they increase their probability of a fatal crash (e.g., Dee, 2001; Hakim et al., 1991; Martinez-Schnell and Zaidi, 1989; Mast, Benson, and Rasmussen, 1999). For example, using state-level data for the period 1982 to 1998, Dee (2001) found that a 10 percent increase in aggregate vehicle miles traveled was associated with a 2 percent increase in teenage motor vehicle fatalities. In the long run, the substitution of larger, less fuel-efficient cars may reduce the fatality risk for the occupants of these vehicles conditional on a crash, but it may also increase the chances of individuals in other cars dying in crashes because larger vehicles are present on the road (Gayer, 2004). Moreover, there are differential behavioral responses. In the context of higher gasoline prices, an individual may take the bus to work but continue to drive to bars during the evening hours, and "careful" drivers may be more or less responsive to higher prices than their less careful neighbors. Clearly, the marginal fatality risk associated with the change in VMT is of central importance.

Thus, in working through the three links, we can establish hypotheses regarding the effect of gasoline prices on motor vehicle fatalities per capita and fatalities per VMT. Given the three structural relationships described above, a decrease in gasoline prices is predicted to unambiguously increase motor vehicle crashes per capita in the short run. For example, lower prices are expected to increase gasoline consumption, which increase the number of VMT, which ultimately increase fatalities per capita. In the long run, we also expect gasoline prices to have a negative effect on motor vehicle fatalities per capita, but the magnitude of this effect relative to the short run is somewhat ambiguous. This result hinges on whether individuals are changing vehicles or changing other behaviors in response to shifts in gasoline prices. For example, a shift to a less fuel-efficient vehicle in response to a decrease in price may attenuate the long-run effect, because the increase in gasoline consumption would not necessarily entail more VMT.[2] Alternatively, long-term behaviors, such as moving farther away from one's workplace in the context of lower gasoline prices, would increase the long-run effect because individuals would increase their VMT and their subsequent risk of a fatal crash.

The short-run relationship between gasoline prices and motor vehicle fatalities per VMT is ambiguous. The direction of the relationship hinges on the marginal fatality risk associated with the increase or decrease in VMT. If higher gasoline prices generally lead to the curtailment of the most risky VMT, then a decrease in fatalities per VMT would be expected. Alternatively, if higher gasoline prices generally lead to the curtailment of equally risky VMT, then an effect of gasoline prices on fatalities per VMT would not be expected. Finally, if higher gasoline prices lead to the curtailment of the least risky VMT, then an increase in fatalities per VMT would be expected. One reason to suspect that the first relationship may hold is that high-risk younger and older drivers may be more sensitive to fluctuations in gasoline prices than middle-aged drivers. The long-run effect of gasoline prices on fatalities per VMT would also depend on the fatality risk associated with the marginal VMT over the long term.

[2] There may also be safety issues associated with the size and mix of vehicles on the road. Given a crash, larger vehicles have been shown to be associated with a lower fatal crash risk for their occupants, but they pose negative externalities for occupants in other vehicles (Gayer, 2004).

Surprisingly, the previous work on the effects of gasoline prices on motor vehicle fatalities has been sparse. Leigh and Wilkinson (1991) used a random effects approach to examine the effects of taxes and prices on fatalities per capita for the period 1976 to 1980. The authors concluded that a 10 percent increase in the state gasoline tax would decrease the fatality rate by 1.8 to 2.0 percent. However, given the fact that state gasoline taxes make up only a fraction (roughly 16 percent over the period 1983–2000) of the overall price of gasoline, the magnitude of the authors' estimates is quite large.[3]

A key empirical issue is whether there is sufficient exogenous variation in gasoline prices to detect a change in the motor vehicle fatality rate. According to the Energy Information Administration (EIA, 2001), the cost of a gallon of gasoline reflects several different components, including the cost of crude oil (46 percent), federal, state, and local taxes (28 percent), refining costs and profits (14 percent), and distribution, marketing, and station costs and profits (12 percent). As Figure 2 illustrates, the real price in 2002 dollars of regular unleaded gasoline ranged from a high of $2.13 in 1983 to a low of $1.13 in 1998. Changes in crude oil prices explained much of the variation over time in retail gasoline prices, such as the upward spike during the Gulf War at the end of 1990. Crude oil prices are determined in the international oil market by changes in oil demand, available supply, and the success of the Organization of Petroleum Exporting Countries (OPEC) in controlling its production. Changes in gasoline taxes have not played a major role in the fluctuation of prices at the national level. The states and the federal government have apparently raised taxes in the aggregate only sufficiently to offset general inflation.

This is not to suggest that there is not variation in gasoline prices across states. In the year 2000, prices varied by 15 to 20 cents per gallon with prices lowest in the Southeast and highest in the West and Midwest. Much of this has to do with proximity to supply, state environmental programs, competition in local markets, regional supply disruptions, differential local operating costs, and state taxes (EIA, 2001). Finally, prices are higher in summer and fall and lower in late winter. Over the period 1983 through 2000, the average price in June was 10 cents higher than in March in 2002 dollars.

Rather than estimate each of the three structural relationships linking gasoline prices and motor vehicle fatalities, this paper presents "reduced form" estimates of the effect of gasoline prices on motor vehicle fatalities per capita and per VMT.[4] We employ a fixed-effects model, which exploits within-state variation in the regressors and outcomes, and as a result, controls for time-variant factors that differ across the states. Thus, we net out cross-sectional differences in gasoline prices and motor vehicle fatalities (e.g., proximity to the Gulf Coast). We also use monthly indicator variables to net out seasonal fluctuations (e.g., higher prices in summer months). Finally, we use annual indicators to net out national temporal trends (e.g., the price of crude oil). Thus, controlling for state-, month-, and year-fixed effects, we exploit within-state variation in price over time, such as shifts in state

[3] Porter (1999) argues that roughly half of state gasoline taxes are passed through to consumers. Thus, a potential upper bound for the gasoline tax-fatality elasticity would be around –0.04 (i.e., 8 percent of –0.5, a high-end estimate of the short-run price elasticity of demand for gasoline). The estimates from the Leigh and Wilkinson study are between 4.5 to 5 times greater than this upper bound.

[4] Cook and Moore (2000) discuss the benefits of the reduced form approach as it pertains to the related literature examining alcohol prices and motor vehicle fatalities. Basically, a reduced form approach is preferred because of the potential for large errors in measurement associated with the intermediate variables and because the chain of argument associated with these variables may be imprecise.

excise taxes[5] and other nonmarket factors,[6] to estimate short-run and long-run implications for fatalities per capita and fatalities per VMT.

DATA

The source of all motor vehicle fatality information within this study is the Fatality Analysis Reporting System (FARS). The FARS, collected by the National Highway Traffic Administration, is a census of all motor vehicle crashes involving a fatality. To be included in this census of crashes, a crash had to involve a motor vehicle traveling on a roadway customarily open to the public and had to result in the death of a person (either the occupant of a vehicle or a nonmotorist) within 30 days of the crash. The FARS contains detailed information on the vehicles, drivers, occupants, and non-occupants involved in the crash. Most analyses of the FARS data construct state-year motor vehicle fatality rates. Importantly, however, we construct state-month fatality rates by using specific monthly fatality counts from the FARS in the numerator, and alternatively, either annual population estimates from the Census Bureau or annual VMT estimates from the Federal Highway Administration in the denominator (see Table 1 for descriptive statistics).[7]

Monthly regular grade, unleaded gasoline prices were obtained from EIA surveys of refiners and retailers/resellers for the period 1983 through 2000. EIA gasoline price data were collected from a sample of about 110 refiners/gas plant operators on the EIA-782A form and a sample of about 2200 retailers/resellers on the EIA-782B form. This EIA price information did not include state or federal gasoline taxes, but this information was obtained from the Federal Highway Administration and added to the price information collected from the EIA.[8] Gasoline price information was missing for 27 state-month cells, which were subsequently omitted from the empirical analyses presented below.[9] Monthly unemployment rate data were obtained from the Bureau of Labor Statistics. Annual per capita income information was obtained from the Bureau of Economic Analysis. Finally, a series of motor vehicle regulations—such as seat-belt laws, speed limit laws, and alcohol-related policies—was collected from various published and unpublished sources.[10]

[5] State gasoline taxes make up a significant portion of the overall price of gasoline, accounting for 14.6 percent of the price of regular unleaded gasoline in 2000. Although the average state tax has remained relatively constant over the 1983–2000 period (Figure 2), some states have made changes. New York State, for example, increased the nominal gasoline tax from 8 cents per gallon in 1989 to 14.4 cents in 1990 to 20.8 cents in 1991.

[6] There are multiple nonmarket sources of within-state variation in gasoline prices. For example, states in the Midwest use a unique reformulated gasoline that is produced using ethanol (a product made from local corn). Few refineries outside the Midwest are prepared to produce this reformulation and tight crude oil supplies in the summer of 2000, for example, caused an increase in prices in the Midwest relative to the rest of the United States. Similarly, California operates its own reformulated gasoline program with more stringent requirements than the Federal clean gasoline mandates. California prices are more variable than other states partly because there are relatively few supply sources of its unique blend outside the state (EIA, 2001; GAO, 2001). Thus, if more than one of the state's refineries experience operating difficulties at the same time, then the supply of gasoline falls and prices increase.

[7] Unfortunately, monthly state population and VMT data are not available.

[8] Importantly, both state and federal gasoline taxes are per gallon and not *ad valorem*.

[9] All results presented within this paper are robust to an alternative model that included interpolated values for the 27 missing cases.

[10] The authors thank Thomas Dee for sharing his motor vehicle regulation data for the purposes of validating the adoption dates of these various policies. See Morrisey and Grabowski (in press) for a discussion of the regulatory data sources.

Table 1. Descriptive statistics: State-month panel data, 1983–2000 (N = 10,341).

Variable	Mean	SD
Traffic fatalities per 1,000,000 individuals	15.66	6.53
Traffic fatalities per 1,000,000, ages 15–17	23.48	19.78
Traffic fatalities per 1,000,000, ages 18–20	26.65	20.77
Traffic fatalities per 1,000,000, ages 21–24	29.39	20.21
Traffic fatalities per 1,000,000, ages 75+	22.46	16.58
Traffic fatalities per one million vehicle miles traveled	1.72	0.69
Real ($2002) gasoline price in cents	149.27	28.51
Illegal per se at 0.08 BAC	0.15	0.35
Illegal per se at 0.10 BAC	0.74	0.44
Administrative license revocation	0.57	0.49
Mandatory seat belt law—primary enforcement	0.17	0.37
Mandatory seat belt law—secondary enforcement	0.52	0.50
65 MPH speed limit	0.47	0.50
70+ MPH speed limit	0.15	0.36
State unemployment rate	5.84	2.16
Real ($2002) state personal income per capita	25,539.92	4,239.92

Note: Alaska, Hawaii, and the District of Columbia are omitted.

EMPIRICAL METHODS

The empirical model exploits the panel nature of the FARS data to examine the effect of monthly gasoline prices on the motor vehicle fatality rate. During the period January 1983 through December 2000, we had access to 10,341 observations (i.e., 48 states times 12 months times 18 years minus 27 missing observations). The basic specification for the empirical results presented here is:

$$F_{smt} = P_{smt}\beta + Z_{smt}\gamma + v_s + k_m + w_t + \varepsilon_{smt} \qquad (1)$$

where F_{smt} refers to the motor vehicle fatality rate in state s in month m of year t, Z_{smt} includes an intercept and a set of exogenous controls, v_s is a state fixed effect, k_m is a month-specific intercept, w_t is a year-specific intercept, and ε_{smt} is the randomly distributed error term. P_{smt} represents the regular unleaded gasoline price in 2002 dollars. The state-fixed effects control for any fixed state-specific omitted variables correlated with the propensity to change the motor vehicle fatality rate. Such variables may include, for example, religious composition in the state and geographic characteristics. The month-fixed effects control for unobserved seasonal variation in the motor vehicle fatality rate such as weather changes or driving patterns.[11] The year dummies control for national trends in motor vehicle fatality rates that may be correlated with changes in gasoline prices such as federal motor vehicle policies and the progress of motor vehicle and road safety technology. Thus, the basic identification strategy implicit in equation (1) purges the unobserved and potentially confounded cross-sectional heterogeneity by relying on the within-state

[11] Importantly, the month-fixed effects control for monthly average facilities across the United States, and do not control for state-specific variation in weather patterns. This may be problematic if both fatalities and gas prices increase in times of inclement winter weather. To test whether this issue was important, we dropped cold weather states during the winter months; our primary results did not change appreciably.

variation in state gasoline prices across months and by using states that did not change their prices as a control for unrelated time-series variation.[12]

Most published estimates of gasoline demand have treated price as an exogenous variable, arguing that variation in prices reflect differences in state excise taxes and other nonmarket factors. However, it may be reasonable to suppose that gasoline prices tend to be influenced by demand, even after controlling for state, year, and month fixed effects. If there exists such a bias in exploiting within-state variation in gasoline prices, then we argue that this bias runs directly against our relationship of interest. For example, a price increase due to an outward shift in the demand for gasoline will increase fatalities. Thus, the estimates presented within this study can be thought of as lower bounds in evaluating the effect of gasoline prices on motor vehicle fatalities.

The exogenous controls introduced in the vector Z_{smt} include both a set of macroeconomic factors and a series of binary indicators representing various state-level regulations. The two macroeconomic factors are the monthly unemployment rate and the annual per capita income level. Several studies have recognized the importance of including these macroeconomic factors within analyses of state motor vehicle fatality rates (e.g., Evans and Graham, 1988; Ruhm, 1996). The analysis also controls for major state laws unrelated to gasoline prices that may be important in explaining changes in motor vehicle fatality rates over the study period. Two binary indicators are included for mandatory seat-belt laws. Seat-belt laws with primary enforcement allow the police to directly cite a motorist for not wearing a belt. Under secondary enforcement of a seat-belt law, a motorist can only be cited for a violation if they are pulled over for some other reason. Seat-belt laws have been shown to reduce traffic fatality rates (e.g., Evans and Graham, 1991). Two other binary indicators identify those states that have increased their rural interstate speed limit to 65 miles per hour or to 70 or more miles per hour. There is inconclusive empirical evidence regarding the system-wide effect of these speed limit laws, but higher speed limits have been shown to increase rural interstate fatalities (e.g., Farmer, Retting, and Lund, 1999; Greenstone, 2002; Lave and Elias, 1997). Finally, the analysis controls for three binary indicators of state laws related to drunk driving.[13] The first indicator equals 1 for those states that make it illegal per se to drive with a BAC of 0.08. The second indicator equals one for those states in which it is illegal per se to drive with a BAC of 0.10. Finally, the third indicator equals one if the state's licensing authority is allowed to suspend driving privileges before any court action related to a charge of drunk driving. There is recent evidence that the use of each of these three policies may be important in lowering the motor vehicle fatality rate (Dee, 2001).

In the overall fatality models, F_{smt} represents the natural logarithm of the monthly motor vehicle fatality rate with either the state population or the number of vehicle miles traveled in the denominator. In modeling the motor vehicle fatality rate, a semi-log specification is frequently employed due to skewness (e.g., Dee, 2001). In

[12] Importantly, there is substantial within-state variation in real gasoline prices and fatalities over time. The mean price measure is $149.27 (in 2002 dollars) and the overall standard deviation (SD) is 28.51 with a between-state SD of 6.95 and a within-state SD of 27.67. The mean fatality rate per 1,000,000 individuals is 15.66 with an overall SD of 6.53 and a between-state SD of 4.55 and a within-state SD of 4.73. The R^2 from a regression of gas prices on the state, year, and month fixed effects is 0.89, implying that there is residual variation to identify the model.

[13] A number of other alcohol-related policies have been considered within the literature. The three policies we include within our model have generally been shown to be important predictors of motor vehicle fatalities (Dee, 2001). As a robustness check, the inclusion of a broader set of alcohol-related policies such as dram shop laws, beer taxes, mandatory jail time, and zero tolerance and minimum drinking age laws within the younger driver specifications did not substantially alter the results presented within this study.

our age-specific fatality models, however, some state-month cells have zero fatalities. To accommodate these zero cases, fatalities are treated as a count on the left-hand side of the model and the natural log of the age-specific population is included on the right-hand side of the model.[14] Conventional count data models do not generate consistent estimates when cross-sectional fixed effects are introduced because of the "incidental parameter" problem (Lancaster, 2000). Thus, following work by Dee and Evans (2001), this analysis employs the conditional maximum likelihood approach for negative binomial models developed by Hausman, Hall, and Griliches (1984). The negative binomial model is less restrictive than a Poisson regression because it accommodates the presence of over-dispersion in the counts. Similar to the semi-log model, the estimates generated by the negative binomial model can be interpreted as the proportionate change in the given motor vehicle fatality count (Cameron and Trivedi, 1998). An alternative method of modeling the dependent variable in both the overall and age-specific cases would have been to treat the state-month fatality rate as grouped data generated by a binary process (e.g., whether an individual dies in a motor vehicle crash in a given month). In a set of sensitivity checks that are available from the authors upon request, the semi-log model estimates presented here proved to be robust in magnitude and precision to grouped logistic models.

The empirical literature on gasoline demand has argued that different short- and long-run effects are associated with price changes (e.g., Baltagi and Griffin, 1984; Dahl and Sterner, 1991). In the context of our study, this observation implies that a change in today's gasoline price may have implications for both today's and tomorrow's motor vehicle fatality rate. To test for any longer-run effects of gasoline price changes, versions of the model specified above were estimated with one-year (P_{smt-1}), two-year (P_{smt-2}), three-year (P_{smt-3}), and four-year (P_{smt-4}) lags. A final methodological point concerns the likely presence of heteroskedasticity in grouped state-level data, which may bias the estimates of the parameter standard errors. For the semi-log estimates, we addressed this potential issue by adjusting the standard errors using the Huber-White (or Sandwich) robust estimator.[15]

RESULTS

Gasoline Prices and Short-run Fatality Rates

The first set of results examines the effect of real gasoline prices on the short-run motor vehicle fatality rates (see Table 2). The first column presents the effect of gasoline prices on motor vehicle fatalities per capita controlling only for the state unemployment rate and the real per capita income and excluding the state-, year-, or month-fixed effects. These potentially confounded cross-sectional estimates indicate that a one-cent increase in the real gasoline price is associated with a 0.06 percent statistically significant decline in monthly fatal crashes per capita. Evaluated at the mean price, this coefficient implies an elasticity of –0.09. The second column introduces state, year, and month fixed effects to the specification in column 1. This result implies an elasticity of –0.14, which is well within the plausible range implied by estimates of the short run price elasticity of demand for gasoline from the literature. The third column presents a fuller specification of the model contained in column 2 with the addition of binary indicators for alcohol, speed limit, and seat

[14] Unfortunately, the number of VMT are not available for different age groups by state and year.
[15] This alternative variance estimator produces consistent standard errors even in cases in which the residuals are not identically distributed.

Table 2. Least squares estimates of a semi-log model for motor vehicle fatalities per capita and per vehicle mile traveled, 1983–2000. (N = 10,341).

	Fatalities per capita			Fatalities per VMT		
Independent Variables	(1)	(2)	(3)	(4)	(5)	(6)
Real gasoline price	–0.00063**	–0.00096**	–0.00108**	–0.00061*	–0.00071*	–0.00059*
	(0.00014)	(0.00028)	(0.00028)	(0.00029)	(0.00028)	(0.00029)
Illegal per se at 0.08 BAC	—	—	–0.021	0.010	–0.040**	0.011
			(0.015)	(0.019)	(0.015)	(0.019)
Illegal per se at 0.10 BAC	—	—	0.016	–0.018	–0.014	–0.024
			(0.011)	(0.013)	(0.011)	(0.013)
Administrative license revocation	—	—	–0.0469**	–0.017	–0.0386**	–0.030**
			(0.0089)	(0.011)	(0.0089)	(0.011)
Primary seat belt law	—	—	–0.060**	–0.035*	–0.055**	–0.027
			(0.13)	(0.015)	(0.013)	(0.015)
Secondary seat belt law	—	—	–0.012	–0.035**	–0.016	–0.026*
			(0.011)	(0.011)	(0.011)	(0.011)
65 MPH speed limit	—	—	–0.029*	–0.018	–0.038**	–0.031*
			(0.012)	(0.013)	(0.012)	(0.013)
70+ MPH speed limit	—	—	0.081**	0.037	0.057**	0.017
			(0.017)	(0.022)	(0.017)	(0.022)
State unemployment rate	–0.036**	–0.0503**	–0.0469**	–0.0390**	–0.0374**	–0.0387**
	(0.017)	(0.0024)	(0.0025)	(0.0031)	(0.0025)	(0.0031)
Real per capita income ($1000s)	–0.06273**	–0.0025	–0.0020	0.0172**	–0.0031	0.0024
	(0.00093)	(0.0036)	(0.0038)	(0.0057)	(0.0038)	(0.0057)
Year fixed effects	No	Yes	Yes	Yes	Yes	Yes
Month fixed effects	No	Yes	Yes	Yes	Yes	Yes
State fixed effects	No	Yes	Yes	Yes	Yes	Yes
State-specific time trends	No	No	No	Yes	No	Yes
R^2	0.30	0.67	0.67	0.69	0.65	0.67

Huber-White standard errors are presented in parentheses. The mean value of the gasoline price variable is 149.27 cents.
* = statistically significant at 5 percent level; ** = statistically significant at 1 percent level.

belt laws. This full model implies a slightly larger elasticity of –0.16. The gasoline price estimates in both columns 2 and 3 are statistically significant. Thus, based on a full model that controls for fixed-effects, macroeconomic factors, and a set of traffic-related laws, a 10 percent decrease in the real gasoline price is associated with a 1.6 percent increase in fatal crashes per capita.

As a potential robustness check to this traditional fixed-effects specification, there may exist unobserved factors changing over time within states that are confounding these results. For example, within-state variation in fatalities may relate to such state-specific factors as the degree of law enforcement or the design and safety of state highways. If changes over time in these unobserved factors are correlated with the within-state variation in both gasoline prices and motor vehicle fatalities, then the traditional fixed effects specification may result in biased estimates. Thus, column 4 introduces both annual (t^*v_s) and seasonal (m^*v_s) state-specific linear time trend variables to the model specified in equation 1 in order to control for unobserved determinants of within-state variation in motor vehicle fatalities.[16] This model indicates that a one-cent increase in the real gasoline price is associated with a statistically significant 0.06 percent decline in the motor vehicle fatality rate. Evaluated at the mean price, this coefficient implies an elasticity of –0.09. Thus, even after controlling for state-specific time trends, we still obtain a strong negative short run relationship between gasoline prices and fatalities per capita.

The final two columns of Table 2 present specifications that examine the effect of gasoline prices on motor vehicle fatalities per VMT. In both specifications, a decrease in the price of gasoline was associated with a statistically significant increase in fatalities per VMT. In a specification that included fixed effects and the full set of covariates (column 5), a 1-cent increase in gasoline prices decreased monthly fatalities per VMT by 0.07 percent. Evaluated at the mean price, this coefficient implies an elasticity of –0.11. In a specification that incorporated annual and seasonal state-specific linear time trends (column 6), the model implied a slightly smaller elasticity of –0.09.

In sum, the short-run effect of gasoline prices on fatality rates is negative and statistically significant, regardless of whether the population or VMT measure was used in the denominator to construct the fatality rate. The robustness of the results across the alternative denominators suggests that, in the context of higher gasoline prices, the foregone VMT are those higher risk miles. Age-specific estimates presented below provide a further test of this supposition.

Gasoline Prices and Long Run Fatality Rates

Table 3 presents eight fixed-effects specifications that incorporate lagged gasoline prices in an effort to examine whether there is a differential long-term effect of gasoline price on motor vehicle fatality rates.[17] The upper panel of the table examines fatalities per capita and the lower panel examines fatalities per VMT. The first column of both panels presents a model that includes a 1-year lagged gasoline price measure and each subsequent column introduces an additional year-lag to the model.

[16] Importantly, these time trends capture only those confounders that trend linearly within states. Additionally, the results are robust to replacing the seasonal and annual trend with a single monthly linear trend (i.e., January 1983 = 1 and December 2000 = 216).

[17] Based on both the adjusted R^2 and Schwarz's criterion (Greene, 2000), the appropriate lag length is less than five. Additionally, in macroeconomic time series models with quarterly data, some studies include both first (one-month) and twelfth (one-year) order lags. Our results presented in Table 3 are robust to the inclusion of a first order lag.

Table 3. Least squares estimates of a semi-log model for motor vehicle fatalities, 1983–2000.

Independent Variables	Fatalities per capita			
	(1)	(2)	(3)	(4)
Real gasoline price	–0.00108**	–0.00091**	–0.00092**	–0.00080*
	(0.00028)	(0.00029)	(0.00030)	(0.00035)
Real gasoline price, 1-year lag	–0.00067*	–0.00060*	–0.00055	–0.00063*
	(0.00027)	(0.00028)	(0.00029)	(0.00030)
Real gasoline price, 2-year lag	—	–0.00043	–0.00044	–0.00043
		(0.00031)	(0.00032)	(0.00032)
Real gasoline price, 3-year lag	—	—	0.00010	0.00011
			(0.00034)	(0.00035)
Real gasoline price, 4-year lag	—	—	—	0.00023
				(0.00030)
Sum of price coefficients	–0.00175	–0.00194	–0.00182	–0.00152
Joint F-test	9.87**	5.36**	3.66**	2.22*
R^2	0.68	0.68	0.68	0.68
N	9,749	9,164	8,581	8,003
	Fatalities per vehicle mile traveled			
	(5)	(6)	(7)	(8)
Real gasoline price	–0.00072*	–0.00061*	–0.00066*	–0.00046
	(0.00029)	(0.00029)	(0.00030)	(0.00035)
Real gasoline price, 1 year lag	–0.00031	–0.00025	–0.00024	–0.00039
	(0.00027)	(0.00028)	(0.00029)	(0.00030)
Real gasoline price, 2 year lag	—	–0.00002	–0.000035	–0.00011
		(0.00031)	(0.00032)	(0.00032)
Real gasoline price, 3 year lag	—	—	0.00043	0.00046
			(0.00033)	(0.00034)
Real gasoline price, 4 year lag	—	—	—	0.00054
				(0.00030)
Sum of price coefficients	–0.00103	–0.00088	–0.00051	0.00004
Joint F-test	3.68*	1.66	1.50	1.44
R^2	0.63	0.63	0.63	0.63
N	9,749	9,164	8,581	8,003

Huber-White standard errors are presented in parentheses. All models include state, year, and month fixed effects and the unemployment rate, the real per capita income and binary indicators for BAC 0.08 law, BAC 0.10 law, administrative license revocation law, primary seat-belt law, secondary seat-belt law, speed limit law of 65 MPH, and speed limit of 70 MPH or greater. The mean value of the gasoline price variable for the full time series is 149.27 cents.

* = statistically significant at 5 percent level; ** = statistically significant at 1 percent level.

In each of the four per capita specifications, the short-run (monthly) elasticity remains relatively stable to the results presented above in Table 2. That is, the implied short-run elasticities range from –0.12 to –0.16. The 1-year and 2-year lag coefficients generally imply a larger long-run effect of gasoline price on motor vehicle fatalities per capita, although the 2-year lag coefficients are not statistically significant in any of the four specifications. The 3-year and 4-year lag coefficients are positive, but are not statistically significant. When the gasoline coefficients are added to each specification to calculate a long-run elasticity of demand, the 1-year elasticity of demand implied by the estimates in specification 1 is –0.26, the 2-year elasticity from specification 2 is –0.29, the 3-year elasticity is –0.27, and finally, the 4-year elasticity is –0.23. The joint F-test of the coefficients is statistically significant in all four cases.

In turning to the lower panel of Table 3, the short-run estimate is statistically significant and relatively stable across the first three specifications, implying an elasticity in the range of –0.09 to –0.11. However, there is only limited support for a long run effect of gasoline prices on fatalities per VMT. The joint F-test of the coefficients is only statistically significant in specification 5, which includes the 1-year lag term. The 1-year elasticity implied by this model is –0.15.

In sum, there is strong long run effect of gasoline prices on fatalities per capita. This finding implies that, in the context of lower gasoline prices, individuals are increasing VMT that result in a greater fatal crash risk. Interestingly however, there was not a larger long-run effect of gasoline prices on fatalities per VMT, implying that the additional VMT were not associated with a marginally higher risk. Although we will return to the implications of these results for aggregate fatality trends below, the short and long run gasoline price estimates are consistent with the per capita and per VMT fatality time trends in Figure 1. That is, lower gasoline prices over the 1990s contributed to stabilization in the per capita fatality trend and a decline in the per VMT trend, although at a rate much slower than in previous decades.

Gasoline Prices and Age-specific Fatalities

Because certain age groups use automobiles largely for leisure-related activities, these cohorts may have a demand for driving that is more price elastic. In particular, younger and older drivers may experience greater changes in driving in response to changes in prices. This also may imply different mortality correlations with price changes. Table 4 presents conditional maximum likelihood estimates of a fixed effect negative binomial model for motor vehicle fatalities for 15- to 17-year-olds, 18- to 20-year-olds, 21- to 24-year-olds, 75-year-olds and older, and all ages. The negative binomial model, rather than the semi-log model utilized above, is employed to account for the presence of month-state cells with zero fatalities for the age-specific categories. The estimates generated by this model can be interpreted as the proportionate change in the given motor vehicle fatality count. Each of the models includes both a contemporaneous gasoline price measure along with 1- and 2-year gasoline price lags. There is basically no effect for the 15- to 17-year-olds and the oldest age category. However, there is a statistically significant and relatively large effect of gasoline prices on the motor vehicle fatality rate for both 18- to 20-year-olds and 21- to 24-year-olds. A 10-cent decrease in the real gasoline price is associated with a 4.4 percent increase in the long-run fatality rate among 18- to 20-year-olds and a 3.9 percent increase among 21- to 24-year-olds. Put alternatively, these result imply elasticities of –0.66 and –0.58, respectively. As a robustness check,

Table 4. Conditional maximum likelihood estimates of a fixed-effect negative binomial model for motor vehicle fatalities by age, 1983–2000 (N = 9164).

Independent Variables	Age 15–17	Age 18–20	Age 21–24	Age 75+	Overall
Real gasoline price	–0.000006	–0.00170**	–0.00219**	–0.00021	–0.00087**
	(0.000575)	(0.00046)	(0.00042)	(0.00052)	(0.00020)
Real gasoline price,	–0.00107	–0.00174**	–0.00144**	–0.00127*	–0.00090**
1-year lag	(0.00061)	(0.00048)	(0.00045)	(0.00053)	(0.00020)
Real gasoline price,	–0.00020	–0.00096*	–0.00027	–0.00040	–0.00051*
2-year lag	(0.00063)	(0.00049)	(0.00045)	(0.00052)	(0.00021)
Illegal per se at 0.08 BAC	–0.018	0.060	0.021	0.020	–0.002
	(0.039)	(0.032)	(0.031)	(0.034)	(0.013)
Illegal per se at 0.10 BAC	–0.022	0.035	0.0004	0.023	0.012
	(0.033)	(0.027)	(0.0251)	(0.029)	(0.010)
Administrative license	–0.018	–0.055**	–0.052**	–0.004	–0.042**
revocation	(0.022)	(0.018)	(0.016)	(0.019)	(0.007)
Primary seat belt law	–0.099**	–0.049	–0.055*	–0.075*	–0.061**
	(0.035)	(0.028)	(0.025)	(0.032)	(0.011)
Secondary seat belt law	–0.009	–0.027	–0.008	–0.030	–0.010
	(0.023)	(0.019)	(0.017)	(0.022)	(0.008)
65 MPH speed limit	–0.013	0.024	–0.036	–0.008	–0.0320**
	(0.026)	(0.020)	(0.019)	(0.022)	(0.0084)
70+ MPH speed limit	0.023	0.047	–0.007	0.067*	0.060**
	(0.040)	(0.032)	(0.030)	(0.033)	(0.013)
State unemployment rate	–0.0614**	–0.0538**	–0.0462**	–0.0324**	–0.0460**
	(0.0063)	(0.0051)	(0.0046)	(0.0057)	(0.0021)
Real per capita income	–0.003	0.0019	0.0150*	0.0055	–0.0027
($1000s)	(0.010)	(0.0082)	(0.0075)	(0.0088)	(0.0034)
Natural log of age-	0.74**	0.743**	0.814**	0.36**	0.724**
specific population	(0.12)	(0.092)	(0.083)	(0.12)	(0.056)
Dependent variable mean	4.73	7.35	8.47	6.00	74.19

Standard errors are presented in parentheses. The mean value of the gasoline price variable for the full time series is 149.27 cents.
* = statistically significant at 5 percent level; ** = statistically significant at 1 percent level.

the overall results presented in the final column are very similar to the results from the semi-log model of the fatality rate presented in column 2 of Table 3. A 10-cent decrease in the real gasoline price is associated with an overall 2.3 percent increase in fatalities over the long run, which implies an elasticity of –0.34. Thus, the long-run fatality effects associated with lower gasoline prices for 18- to 24-year-olds are almost twice as large as for individuals of all ages.

The results the in final column of Table 4 also provide an opportunity to summarize the other variables included within the model and compare the relative magnitude of the effects for the regulatory policies to the long run gasoline price effect. Additionally, it is useful to benchmark our results against previous work in the motor vehicle literature. The alcohol policy measures tell a mixed story. Both of the BAC indicators are associated with small and statistically insignificant effects, but administrative license revocation is associated with a statistically significant 4.2 percent decline in motor vehicle fatalities. Recent work by Dee (2001) found a statistically significant negative effect of all three of these variables on the annual motor vehicle fatality rate. The seat-belt findings were in the expected direction. Primary seat-belt laws were associated with a 6.1 percent decline in motor vehicle fatalities and secondary seat-belt laws were associated with a 1.0 percent decline, although only the primary seat-belt law result was statistically significant. In earlier work, Evans and Graham (1991) found that the introduction of seat-belt laws reduced annual fatalities by 8 percent. The finding for the speed limit laws are somewhat paradoxical in that the implementation of a 65 MPH speed limit was associated with a 3.2 percent decline in fatalities while a 70+ MPH speed limit was associated with a 6 percent increase in motor vehicle fatalities. The existing literature with respect to these policies has been inconclusive regarding their effect on overall motor vehicle fatalities (Grabowski and Morrisey, 2001).

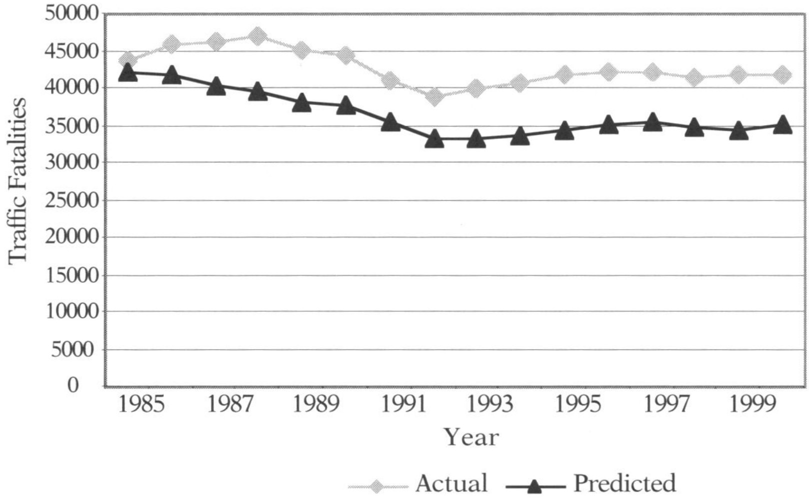
In comparing the size of the gasoline price findings with the regulatory policies, it is clear that gasoline prices are a relatively important determinant of motor vehicle fatalities. The magnitudes of the fatality effects associated with an administrative license revocation law (–4.2 percent), primary seat-belt law (–6.1 percent), and the 70+ MPH speed limit law (6.0 percent) are roughly comparable to a 20-cent decline in the gasoline price sustained over the long run (4.6 percent). A 20-cent shift is a plausible change given the observed volatility in gasoline prices over time. Consider the period of September and October in 2001 when national gasoline prices fell on average 29 cents in a 6-week period and prices in the Midwest fell 57 cents over an 8-week period.

DISCUSSION

One interpretation of the stability of motor vehicle fatality rates over the 1990s is that recent policies and innovations have been ineffective relative to historic initiatives. This paper offers an alternative explanation that these rates have stabilized partly due to gasoline prices. Toward this end, the model estimates can be used to project the number of fatalities assuming the real price of gasoline had remained constant at its 1983 level (see Figure 3). Based on the coefficient estimates from the final column of Table 4, if prices had remained constant at $2.13 (in 2002 dollars) over the period 1983 through 2000, then over 92,000 lives would have been saved between 1985 and 2000. To put this number into some context, approximately 42,000 individuals died annually in motor vehicle crashes over the period 1995 through 2001. Thus, the total number of lives saved over the 16-year period would be comparable to having no motor vehicle related deaths for just over 2 years.

As an alternative benchmark, the magnitude of the long-run change in fatality rates from gasoline price changes is on a par with many of the policy interventions that states have implemented to try to reduce motor vehicle fatality rates. For example, the National Academy of Sciences concluded that the National Maximum Speed Limit of 55 MPH initially saved between 2000 and 4000 lives annually (NRC, 1984).[18] Using the upper bound estimate, a 55 MPH speed limit would have saved 72,000 lives over the period 1983 through 2000, which is 20,000 fewer than our estimate of the number of lives saved had the price of gasoline remained constant at its 1983 level. Clearly, gasoline prices have had a meaningful impact on motor vehicle fatalities in the United States over the last two decades. Moreover, a partial expla-

[18] Our estimates and other recent work examining speed limit laws and motor vehicle fatalities have not always substantiated the large safety benefits found in the National Research Council (NRC, 1984) report. See Grabowski and Morrisey (2001) for a review of the recent literature.



Note: Fatalities are predicted using the current and lagged gasoline price coefficient estimates from the final column of Table 4.

Figure 3. Actual versus predicted traffic fatalities assuming 1983 gasoline prices (1985–2000).

nation for the relative stability of the vehicular fatality rate in the United States over the last decade is that some of the effects of policy interventions such as seat-belt laws and "driving under the influence" (DUI) policies have been attenuated by the trend toward lower real gasoline prices over the period.

These findings may suggest that gasoline taxes should be increased to reduce motor vehicle fatalities, especially in light of the fact that the United States has one of the lowest gasoline tax rates in the industrialized world (Porter, 1999). For example, the Italian gas tax rate is 300 percent of the pretax retail gasoline price, and Canada, Australia, and New Zealand's tax rates are roughly 100 percent of the pretax rate. The United States is an outlier of sorts with a gas tax rate (federal plus state and local taxes) roughly 40 percent as large as the pretax rate. Based on our findings, a tax that yields a real increase in the price of gasoline, sustained over time, will reduce fatalities substantially. Moreover, it will have a disproportionate effect in reducing vehicular deaths among young adults. However, more than auto safety must be considered in such a policy decision. Higher gasoline taxes entail deadweight costs in that drivers will give up some travel that is currently worth more than the cost of travel. The policy issue is whether the (statistical) lives saved are worth more than the value of the travel foregone, an issue beyond the scope of this paper.

In spite of the potential dead-weight costs associated with higher gasoline taxes, there is some evidence that both the federal government and a number of states are currently considering dramatic increases in gasoline taxes (Anderson, 2003). At the federal level, there is a proposal to phase in a series of annual increases that would raise the federal gasoline tax from the current 18.4 cents per gallon to more than 33 cents by 2009. At the state level, budget shortfalls due to the recent economic recession was partially responsible for 5 states passing tax hikes in 2002 with an additional 26 states currently considering tax increases. If the various federal and state proposals become law, they could increase the average price of gasoline by 25 cents per gallon over the next several years. If these increases occur, the results of this paper suggest there will be important implications for motor vehicle safety.

This paper has shown that traffic safety officials must think broadly regarding the interpretation of aggregate motor vehicle fatality trends. We argue that declining gasoline prices are an important reason why fatality rates have largely leveled off over the 1990s. Based on this study, it would be a mistake to conclude from aggregate trend data alone that traffic laws related to drinking and safety belts and innovations such as airbags have not been successful in saving lives over the past decade.

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