The Asymmetric Impact of Oil Prices on Gasoline Prices

Dale Bremmer*
Department of Humanities and Social Sciences, Rose-Hulman Institute of Technology

Kevin Christ**
Department of Humanities and Social Sciences, Rose-Hulman Institute of Technology

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* Department of Humanities and Social Sciences, Rose-Hulman Institute of Technology, 5500 Wabash Avenue, Terre Haute, IN 47803. Tel. 812-877-8456. Fax: 812-877-8909. E-mail: dale.bremmer@rose-hulman.edu

** Department of Humanities and Social Sciences, Rose-Hulman Institute of Technology, 5500 Wabash Avenue, Terre Haute, IN 47803. Tel. 812-872-6226. Fax: 812-877-8909. E-mail: kevin.christ@rose-hulman.edu

I. Introduction

After a lengthy period of fairly stable gasoline prices, recent volatility has renewed an old debate over the nature of retail gasoline price adjustment to shocks in the cost of crude oil. It is not uncommon to hear people claim that retail gasoline prices rise faster than they fall in response to changes in the world price of crude oil, or even that increases in crude oil prices are passed on in the form of higher gasoline prices while decreases in crude oil are not. This commonly held view that gasoline price adjustment is asymmetric thus typically has two dimensions -- speed and magnitude. Taken literally, the second of these claims is clearly implausible, as it would lead to ever-widening cost-price margins. That cost-price margins have not exhibited a long-term secular increase does not, however, negate the possibility of a meaningful magnitude effect for consumers. If cost increases are passed on more quickly than cost decreases, then a contemporaneous magnitude effect will arise. In this paper, we investigate whether the magnitude of price adjustment is indeed asymmetric, and if so, whether there are notable differences in the effect for different regions of the country and different frequency of the data - - monthly, weekly, or daily.

The period from mid-summer to late fall 2001 illustrates why so many people believe that gasoline price adjustments are asymmetric. For the week of July 16, crude oil was 54.1 cents per gallon and the average retail price of regular gasoline was $1.33.[1] Over the next eight weeks, the prices of both rose, so that by the week of September 10, crude oil was 60.5 cents per gallon and the retail price of gasoline peaked at $1.51. Thus, a 12% increase in crude oil prices was associated with an almost 14% increase in the price of gasoline. From this retail price peak, the price of both crude oil and gasoline fell over the following eight weeks, so that by the week of November 5, crude oil was 41.4 cents per gallon and the retail price of gasoline was $1.17. Thus, a 32% decrease in the price of crude oil was associated with only a 23% fall in the price of gasoline. Thus, while the magnitude of the retail price increase closely mirrored the late-summer increase in the price of crude oil, the autumn decline in crude oil prices was only partially reflected in falling retail gasoline prices. Such episodes provide anecdotal evidence of asymmetric responses that are unfavorable for consumers, and typically give rise to charges of market power abuse.

The difficulty with such analysis is that it is easy to find historical episodes that illustrate different relationships. In the five weeks from early December 1998 to mid January 1999, the price of crude oil rose 19%. In the ensuing four weeks up to early February, the price of crude oil fell almost 13%. During this time, gasoline prices barely move from a range of \$0.91 to \$0.94. In another episode, from mid-June to the end of July in 1998, the price of crude oil rose by 19%, but retail gasoline prices actually fell by about 1.4%.

Asymmetric output price adjustment to changes in input costs has been called a "stylized fact" for many markets (Peltzman, 2000), one that poses significant challenges to prevailing theories of price determination. Most models of price determination allow no distinction for different kinds of input price shocks on output prices -- equal magnitude increases and decreases in input prices should result in symmetric changes in output prices. It is often assumed that asymmetric price responses of the type investigated here are the result of market power being exercised at different points in the supply chain. Speculation about the causes of asymmetries is not limited to monopoly power, however, and includes focal point pricing, inventory and menu costs, and signal extraction problems brought on by uncertainty (Borenstien, et al, 1997). While of continuing theoretical interest, such speculation is not the focus of this paper. Instead, we simply investigate the temporal and regional nature of asymmetric responses, if they exist at all.

Gasoline markets would seem to offer a classic opportunity for this phenomenon to emerge, as there is such a clear relationship between the input and the output. Previous studies

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[1] All prices are derived from weekly data available from the Energy Information Administration on its web site.

seem to concur that retail gasoline prices respond more quickly to input cost increases than to input cost decreases, although the timing and significance of the effect is open to question. While the hypothesis of symmetry in the overall response of retail prices to increases or decreases in the price of crude oil seems supportable, the pattern of the adjustment does seem to be different, thus giving rise to short-term magnitude effects (Karrenbrock, 1991). These different magnitude effects seems to prevail for about six to eight weeks (Borenstein, et al, 1997), after which the distinction between an input price increase and decrease seems to fade.

Table 1 presents simple evidence of this time structure to the adjustment process. The contemporaneous (week-on-week) correlation between changes in the price of crude oil and regular gasoline is positive but weak. As the time horizon lengthens, the correlation between changes in crude oil prices and retail gasoline prices becomes stronger.

Attention to the potentially asymmetric price adjustment seems to increase with the volatility of crude oil prices. During periods of relatively stable crude oil prices, retail gasoline prices also exhibit relative stability. But crude oil price spikes heighten consumers' awareness of the sensitivity of travel costs to the vagaries of crude oil prices. Recent years have seen highly volatile crude and gasoline prices. Table 2 provides a very broad summary of gasoline and crude oil price behavior over recent years. Two features stand out regarding gasoline prices. First is the changing nature of the medium-term trends in prices. A period of relative price stability for most of the early 90s was followed by a 2-year period of declining prices, which in turn has been followed by the most recent period of rising prices. The second feature is the relative volatility of these price changes. Compared to the early 1990s, prices have been more volatile since 1997, and much more so since 1999. This volatility has introduced new levels of uncertainty into gasoline prices.

To investigate fully the existence and nature of pricing asymmetry, we use price data from the Energy Information Administration, which maintains rich sets of daily, weekly and monthly time series for gasoline and crude oil prices. In this paper, we are interested mainly in furthering our understanding of the dimensions of any asymmetry over time and across regional markets. While these data do not account for differences in sales and excise taxes across states, we don't feel these effects are significant to our results, or severely limit the usefulness of our regional analysis.

II. The regression model and its data

The specification of the general regression model

We employ a simple regression model to test for a contemporaneous, asymmetric response of retail gasoline prices to changes in the price of crude. Let P_t^g denote the retail price of regular gasoline in time t, while P_t^c is the price of crude in time t. The price of gasoline is affected by supply-side factors such as the price of crude and demand-side factors such as household income and whether it's the summer vacation season. The supply of gasoline and its price is also influenced by the price of alternative fuels derived from crude oil such as jet fuel and heating oil. Let all the other relevant determinants of the price of gasoline, other than the price of crude, be denoted by $\mathbf{x_t}$, a k x 1 column vector of the other explanatory variables at time t. Let X_{it} be the ith element of $\mathbf{x_t}$, or in other words, X_{it} is one of the other k explanatory variables, other than the price of crude, that affects the price of regular gasoline at time t. To identify times when the price of crude increased, let D_t be a dummy variable such that D_t equals 1 if P_t^c - $P_{t-1}^c > 0$, or D_t equals 0 if P_t^c - $P_{t-1}^c \leq 0$.

The general specification of the regression model is

$$P_t^g = \beta_0 + \beta_1 P_t^c + \beta_2 D_t \cdot P_t^c + \sum_{i=1}^{k} \alpha_i X_{it} + u_t \qquad (1)$$

where â$_0$, â$_1$, â$_2$, and vector **á** are the unknown regression parameters to be estimated, while u_t is the error term of white noise. If crude oil prices increase from time t-1 to time t, then D_t equals 1 and equation (1) becomes

$$P_t^g = \beta_0 + (\beta_1 + \beta_2)P_t^c + \sum_{i=1}^{k} \alpha_i X_{it} + u_t \, . \tag{2}$$

Conversely, if the price of crude remains constant between periods t-1 and t, or if the price of crude declines between times t-1 and t, then equation (1) becomes

$$P_t^g = \beta_0 + \beta_1 P_t^c + \sum_{i=1}^{k} \alpha_i X_{it} + u_t \, . \tag{3}$$

From equation (2), it follows that if the price of oil increases between periods t-1 and t then

$$\frac{\partial P_t^g}{\partial P_t^c} = \beta_1 + \beta_2 \, ; \tag{4}$$

however, according to equation (3), if the price of oil remains constant or falls between periods t-1 and t, then

$$\frac{\partial P_t^g}{\partial P_t^c} = \beta_1 \quad . \tag{5}$$

Thus, the statistical test to determine the presence of a contemporaneous, asymmetric price response simplifies to a simple t-test on whether â$_2$ is statistically different from zero. If the estimated coefficient of â$_2$ is statistically significant, then there is statistical evidence of an asymmetric response of the price of gasoline to the price of crude oil. Rejection of the null hypothesis that $\beta_2 \leq 0$ would imply that the price of gasoline responds disproportionately with increases in the price of crude compared to decreases in the price of crude.

In this paper, we test variations of this general specification using data sets with different frequency and different cross-sections. The regression is estimated separately using monthly, weekly, and daily spot prices. The different cross-sections include U.S. national data, regional data, data from five states, and data from six cities.

<u>Monthly data: The regression specification and data</u>

To test for asymmetry with monthly data, we estimate the following model:

$$P_t^g = \beta_0 + \beta_1 P_t^c + \beta_2 D_t \cdot P_t^c + \alpha_1 Y_t + \sum_{i=2}^{12} \alpha_i M_{it} + u_t \qquad (6)$$

where P_t^g, P_t^c, and D_t are as defined above, Y_t denotes monthly disposable nominal income, and M_{it}, i = 2, . . . , 12, are monthly dummy variables used to control for seasonal price effects. The price of gasoline should be directly related to the price of crude; therefore, $\hat{a}_1$ is expected to be positive. Gasoline should be a normal good implying increases in monthly income results in increased demand for gasoline and higher gas prices. Thus, $\acute{a}_1$ is also expected to be positive. Finally, if the price response of gasoline prices is asymmetric to changes in the price of crude, the popular assumption is that $\hat{a}_2$ is positive, implying that the absolute value of the change in gasoline prices is greater when the price of crude increases compared to a decrease in the price of crude.

In the monthly data set, the price of regular gasoline (in cents/gallon) is the U.S. city average series of retail prices for regular gasoline obtained from the Department of Energy (DOE). The price of crude is measured by the U.S. refiner's acquisition cost of crude oil, a composite of both domestic and imported oil, which is measured in dollars per barrel.[2] Data on nominal, monthly disposable personal income is measured in trillions, and this data is obtained from FRED, the time-series, internet database of the Federal Reserve Bank of the St. Louis.[3] The monthly data deals with only national U.S. data, and no other cross-sections are possible with this data.

[2] See the Department of Energy's web site http://www.eia.doe.gov/emeu/mer/prices.html lists both the gasoline price and the refiner acquisition cost of crude in its Monthly Energy Time Series spreadsheets.
[3] FRED's URL is http://www.stls.frb.org/fred/.

Weekly data: The regression specification and data

With weekly data, choosing a proxy variable to capture changes in demand becomes problematic. For monthly U.S. data, disposable personal income clearly captures some of the influences that affect the demand for gasoline. There are fewer proxies to capture demand-side effects with weekly data for both the entire U.S. and particular regions of the U.S. For national data, weekly measures of the money supply captured changes in gasoline demand with mixed success. In addition, weekly dummy variables lack the easy interpretation and statistical significance of monthly dummy variables. Consequently, when dealing with weekly or daily prices, the only explanatory variable considered was the price of crude. The general specification of the regression becomes

$$P_t^g = \hat{a}_0 + \hat{a}_1 P_t^c + \hat{a}_2 D_t \times P_t^c + u_t \quad . \qquad (7)$$

With the weekly data set, P_t^g is the weekly retail price of regular gasoline measured in cents per gallon, while P_t^c is the weekly import weighted price per barrel of crude oil (measured in dollars$ per barrel).[4] Again, as with the monthly data, positive values for $\hat{a}_1$ and $\hat{a}_2$ are expected.

The Department of Energy has several different data sets of weekly time series for the price of regular gas. There are national data consisting of retail prices for the entire country, as well as data for five multi-state regions: the East region, the Midwest region, the Gulf region, the Rockies region, and the West region.[5] The East coast region can be further divided into three smaller multi-state areas consisting of the New England region, the Central Atlantic region, and the Lower Atlantic region. There are data for five states: California, Colorado, Minnesota, New York, and Texas. Over the same time period, the Department of Energy has time-series data for retail regular gas prices in six cities: Chicago, Denver, Houston, Los Angeles, New York City,

[4] http://www.eia.doe.gov/oil_gas/petroleum/data_publications/wrgp/mogas_history.html is the Department of Energy's web cite for historical weekly retail gasoline prices.
[5] The states belonging to each region are listed in Appendix 1.

and San Francisco. These data sets allow testing whether there is an asymmetric price response of the product price to an increase in costs in smaller geographic markets compared to national data.

Daily data: The regression specification and data

Finally, daily spot prices are used to determined whether increases in gasoline prices, caused by an increase in the price of crude, are greater than decreases in gasoline prices when the price of oil falls. Daily spot prices of regular gas and crude oil are obtained from the web pages of the Department of Energy. Due to the lack of good proxies to capture daily demand, the regression model focuses solely on the effect of crude oil prices on the price of gasoline. In a regression identical in structure to equation (7), P_t^g is set equal to the New York spot price of conventional regular gasoline (in cents per gallon), while P_t^c, the price of crude, is measured by the spot price of West Texas Intermediate crude oil (in dollars per barrel).[6]

Econometric issues

When analyzing time series such as gasoline prices, crude oil prices, and monthly income, the problems associated with autocorrelation arise when using time series data. The presence of first-order autocorrelation in the regression specifications listed in equations (6) and (7) is tested using the Bruesch and Godfrey test for serial correlation. If autocorrelation is present, then the regression parameters and ñ, the first-order autocorrelation coefficient are estimated using nonlinear least squares - - an approach which gives estimates that are both asymptotically equivalent to maximum likelihood estimates and asymptotically efficient.[7]

The problems associated with spurious regression results are another econometric problem that may be encountered with time-series data. Studies using macroeconomic time

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[6] The spot prices for crude are found at http://www.eia.doe.gov/oil_gas/petroleum/info_glance/crudeoil.html, the Department of Energy's web site for crude oil data. Spot prices for regular gasoline are from the Department of Energy's web page for gasoline data, http://www.eia.doe.gov/oil_gas/petroleum/info_glance/gasoline.html.
[7] See Davidson and MacKinnon (1993, 329-341).

series have shown when a nonstationary independent variable is regressed on a nonstationary explanatory variable, spurious regression results may occur.[8] In this case, the regression is characterized by high a R^2 and simple, independent t-tests incorrectly reject the hypothesis that the estimated parameters are equal to zero while there is no true underlying relationship between the explanatory variable and the dependent variable. Augmented Dickey-Fuller tests are used to determine whether each data series exhibits a unit root and is, consequently, nonstationary. If the price of regular gas and one of its independent variables are nonstationary, then equations (6) and (7) are estimated using first differences of the variables instead of the variables in level form.

III. Empirical Results

Tables 3 through 11 list the results of the unit root tests and the estimated parameters of equations (6) and (7). In general, the statistical evidence indicates that the prices of regular gas and crude oil are nonstationary, and the absolute change in gas prices is greater with increases, rather than decreases, in the price of crude oil.

Monthly data

Using monthly data, Table 3 reports the unit root tests for monthly income and the prices of regular gasoline and crude oil. In this and all other augmented Dickey-Fuller tests performed in the paper, lags of the first-differences of the data were added until serial correlation was removed. Using level data, the hypothesis that the regular gas price series has a unit root can only be rejected at the ten percent level. The level data of crude oil prices is nonstationary as the hypothesis of a unit root is never rejected. Using first-differenced data, the augmented Dickey-Fuller tests indicate all these series are stationary. Given the weak statistical evidence that the level data of regular gas prices is stationary, the null hypothesis of a unit root is only rejected at the 10% percent level - - the weakest level of statistical significance, these variables were used to

[8] See Granger and Newbold (1974), Nelson and Plosser (1982), and Stock and Watson (1988).

estimate equation (6) both in level and first-differenced forms. The regression results are listed in Table 4.

The first two columns of regression results in Table 4 pertain to level data, while the third and fourth columns in Table 4 refer to the results obtained from data that was first-differenced. Examining the level data results first, the low Durbin-Watson statistic of 0.404 agrees with the results of the Bruesch-Godfrey tests that indicate first-order autocorrelation is present. Correcting for autocorrelation, the results in the second column indicate that the price of regular gas is directly related both to the price of crude and monthly income. The coefficients for both of these variables (2.416 and 7.059, respectively) are both positive and statistically different from zero at the one percent level. Dummy variables for the months of April through November are positive and statistically significant at either the one- or five-percent level. The most surprising result is that the estimated parameter $\hat{a}_2$ is negative and significant. The interpretation here is that a \$1 decrease in price of a barrel of crude causes gasoline prices to fall by 2.416 cents per gallon, but a \$1 increase in the price of crude per barrel causes the price of regular gas to increase only by 2.354 cents per gallon. The results indicate that the response of gasoline prices to changes in crude prices is asymmetric, but the result is opposite the one heard in the press. Here gasoline prices are more responsive to a fall, not an increase, of crude oil prices.

However, not much confidence should be placed on this result because both the price of regular gas and crude oil were shown to be nonstationary. Consequently, the regression results may be spurious and inferences based on standard t-test and F-tests are misleading.[9]

Given the problems of unit roots and spurious regressions, the third and fourth columns of Table 4 refer to results obtained from using first-differenced data. Similar to the results

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[9] Referring to the first column of regression results in Table 4 the R^2 was 0.956 and the Durbin-Watson statistic was 0.404. Granger and Newbold (1974) have argued that, if the value of a regression's R^2 exceeds the value of its Durbin-Watson statistic, there is evidence that the estimation results suffer from the complications caused by spurious regression.

obtained from using level data, the results in the third column indicate autocorrelation is present (the Durbin-Watson statistic was 1.620, and the Bruesch-Godfrey confirmed the presence of AR(1) errors). The results in the fourth column of Table 1 indicate that changes in crude oil are directly related to changes in the price of regular gas (the coefficient of 1.900 is statistically significant at the one-percent level). The positive and statistically significant estimates for the dummy variables on the months of March, April, May, and June indicate that these months have significantly higher priced gasoline than the other months. In a surprising result, the estimated coefficient associated with changes in income, while positive as expected, is statistically insignificant (the t-value is less than 0.05).

Just as surprising is the result that the estimated parameter for $\hat{a}_2$ is statistically insignificant. The estimated parameter of 0.151 is positive, but the associated t-statistic is less than 0.5, indicating that the parameter estimate of 0.151 is not statistically different from zero. Consequently, using monthly data, after taking first differences and correcting for autocorrelation, the results of the fourth column in Table 4 indicate there is no contemporaneous asymmetry between the price of regular gasoline and the price of crude. The estimated value for $\hat{a}_1$ indicates that a $\pm$\$1 change in the price of a barrel of crude implies a $\pm$1.9 cent change in price of a gallon of regular gas.

<u>The results using weekly data of U.S. and regional data</u>

Using the various samples of weekly data, the results of the unit root test are listed in Table 5, while the estimation results of equation (7) are listed in Tables 6 and 7. The regression results from using data in levels is shown in Table 6, while the results in Table 7 come from data that has been first-differenced.

The Dickey-Fuller test results in Table 5 indicate that most of the weekly data series exhibit units roots and are nonstationary. Using level data, only in the sample composed of Midwest gasoline prices was the null hypothesis of a unit root is rejected at the five-percent level (the test statistic is 2.892). Referring to state level data, the hypothesis of a unit root is rejected at the ten-percent level for both Minnesota and New York (the test statistics were 2.855 and 2.739, respectively). Only in the case of Chicago was the hypothesis of unit roots rejected at the one-percent level for samples consisting of level data. Nonetheless, the price of crude was found to be nonstationary, and the all of the series were stationary in first differences.

The results of the levels data in Table 6 mimic the results found for level data in Table 4. Using weekly data and either U.S. or regional data, the estimate for $\hat{a}_2$ is always negative, and, with the exception of the Midwest sample, always statistically significant at the one percent level. Again, this indicates the response of regular gasoline prices to changes in the price of crude is less when crude oil prices increase compared to a decrease in the price of crude. Again, given the low Durbin-Watson of the regressions when there was no correction for autocorrelation and the high statistical significance of ñ when the regressions were corrected for autocorrelation, serial correlation appears to be present. Finally, the regression results again appear to be spurious as values of R^2 range from 0.984 to 0.995, while the Durbin-Watson statistic ranges from 0.10 to 0.17.

Unlike the results obtained from monthly data, once first differences of the data are taken and the regressions are corrected for autocorrelation, Table 7 does contain statistical evidence of contemporaneous price asymmetry of regular gas prices with crude oil prices. This is the “expected” outcome where larger increases in regular gas prices with occur with increases in the price of crude oil increases and smaller decreases in gas prices occur with a decrease in oil prices. The estimate for $\hat{a}_2$, the regression parameter which indicates evidence of an asymmetric response is positive and statistically significant for samples that include U.S. data, East Coast data, both Central and Lower Atlantic data, and the Gulf Coast data. The coefficients for $\hat{a}_2$ in the Rockies and Midwest areas are positive, but insignificant. Regarding the Midwest region, the lack of strong statistical evidence of an asymmetric response is interesting given the allegations of price gauging that were common in that area two summers ago. When corrected for autocorrelation, the estimate for $\hat{a}_2$ for the West Coast data equals –0.081, but it is not statistically significant.

The U.S. results indicate that a $1 increase in the per barrel price of crude causes the price of a gallon of regular gas to increase by 1.3 cents. However, a $1 fall in the price of crude is only accompanied by a decline in the price of gas of 0.611 cents per gallon. These results are similar in magnitude and direction to most results in previous studies. The results in Table 7 also indicate that the price response is larger for U.S. data than it is for state data. For example, the sum of $\hat{a}_1$ and $\hat{a}_2$ is estimated to be 1.343 for the U.S., while it equals 0.822 for the East Coast, 0.458 for New England, 0.626 for the Central Atlantic region and 1.045 for the Lower Atlantic region.

<u>Weekly prices: The state and city samples</u>

Tables 8 and 9 list the results of the weekly data pertaining to cross-sections of selected states and cities. The regression results reported in Table 8 deal with both data in level form and in first differences, but there is no correction for autocorrelation. Table 9 lists the regression results after correcting for first-order serial correlation for both data samples - - data in levels and data in first differences.

Again, the results in Tables 8 and 9 mimic those of Table 4. When using levels data, the estimated value of $\hat{a}_2$ is consistently negative, and generally statistically significant. Once again, this leads to the conclusion that gasoline prices increase less given an increase in the price of oil compared to the decline in gasoline prices that occurs when crude prices fall. All the regressions using data in level form in Table 8 exhibit autocorrelation as the Durbin-Watson statistic ranges from 0.10 to 0.21. The data in levels form in Table 8 also tend to give spurious results as indicated by the relatively high R^2s and the low Durbin-Watson statistics.

The right-hand column of Table 9 lists the regression results after taking first differences and correcting for autocorrelation. The results are mixed and not as strong as previous samples with larger geographic areas. The estimated values of $\hat{a}_1$ are generally positive, as only the estimate for Houston was negative. However, only the positive estimates for California and Minnesota are statistically significant. The estimates for $\hat{a}_2$ reported in the right-hand column of Table 7 tend to be positive as negative estimates are reported only for California and San Francisco. The evidence of price asymmetry is not conclusive as only Houston and Texas had positive and significant estimates. The Texas coefficient was significant only at the five-percent level and the Houston coefficient was significant only at the ten-percent level.

It is interesting that the state and city data do not exhibit any statistically stronger evidence of price asymmetry. If the ability to raise prices disproportionately higher because of an input shock is a characteristic of market power, one would intuitively suspect strong evidence of market power in smaller geographic markets. The data for Chicago shows no strong statistical support for price asymmetry despite allegations of price gauging during the supply restrictions that occurred in the Midwest two summers ago.

<u>Results using daily spot prices</u>

The unit root tests for the spot price for regular gas prices and crude oil prices are listed in Table 10. In a result that is counter to both the monthly and weekly data, the null hypothesis of a unit root is rejected at the one-percent level for each of these data series. Both of these data series are stationary in level form.

Despite both the dependent and independent variables appear to be stationary, the regression results using data in levels form appear to suffer from the problems caused by spurious regression results. In an outcome that is consistent with previous results, the estimation results in the first regression reported in Table 9 are characterized by a high R^2 of 0.86, but a low Durbin-Watson of 0.058. Once again, the estimated value of the coefficient that captures an asymmetric price response, $\hat{a}_2$, is negative, but it is not statistically different from zero. However, when the model is estimated with data in first-differenced form and corrected for serial correlation, the estimated for $\hat{a}_2$ is 0.192, both positive and statistically significant at the five-percent level. These results indicate the presence of an asymmetric response of the product's price to a supply shock. A one dollar increase in the price of a barrel of crude leads to a 2.014 cent increase in the price of regular gas. Since there are 42 gallons of crude in a barrel, a one dollar increase in the price of crude implies a 2.38 cent increase in the price of crude per gallon. These estimates imply an increase in the price of a gallon of crude of 2.38 cents leads to a 2.014 cent price increase in a gallon of gasoline. This almost one-for-one price increase exceeds the estimates of previous studies and it is higher than the estimates from the weekly data reported in Table 7.

The estimates in Table 11 also indicate that when crude prices fall by one dollar per barrel, the price of a gallon of regular gasoline falls by 1.829 cents per gallon. While the difference between 2.014 and 1.829 is statistically significant, one has to wonder whether the difference is economically significant to consumers and firms.

IV. Conclusions

The attempt to determine whether there is a contemporaneous, asymmetric response of regular gas prices to increases and decreases in the price of crude brought mixed results. Regressions using weekly and daily U.S. data provide statistical evidence that the increase in current gas prices given a one dollar per barrel increase in crude prices is greater in absolute value than the decrease in gas prices caused by a one dollar per barrel decrease in crude prices. The estimated response to increases in crude price was larger for daily data than it was for weekly data. Regressions with monthly data did not provide statistically significant evidence of an asymmetric response. With asymmetric behavior being exhibited with daily or weekly data, but not with monthly data, this suggests that if asymmetries exist, they are transitory and short-lived. This confirms the conclusion of previous research.

Weekly data of multi-state U.S. regions also showed credible statistical evidence of asymmetric price responses. The Midwest and West Coast regions where the only regions that the estimated value of the regression coefficient that indicates an asymmetric response was not statistically different from zero. The regression results for the samples pertaining to specific states or cities show little evidence of asymmetric price behavior. If asymmetric pricing in response to changes in input prices is characteristic of market power, and if market power is more likely to exist in smaller geographical markets, it is surprising that that there was no credible evidence of asymmetric price behavior in these smaller markets.

Several thrusts for future research come from the paper. Frequency of the data and whether asymmetric price responses occur in annual, quarterly, monthly, weekly, or daily data needs to be explained. The effect of the market size - - city, state, regional, or national - - on the detection of asymmetric price behavior needs explanation. Model dynamics needs to be explored and maybe additional lagged price changes needs to be introduced. Nonetheless, the results of this paper are interesting and important. Studies using both weekly and daily data indicate that there is an contemporaneous, asymmetric response of gas prices to changes in the price of crude. If the price of a barrel of crude oil increases by $1 per barrel - - an increase in the price of almost 2.38 cents per gallon of crude, the price of regular gas increases between 1.343 and 2.014 cents per gallon. One the other hand, if the price of crude falls by $1 per barrel, or equivalently, 2.38 cents per gallon, the price of crude will fall between 0.611 and 1.812 cents per gallon. Future research will have answer why these responses vary in magnitude across smaller regions.

Table 1
Correlation Between Crude Oil and Gasoline Price Changes
Weekly Percentage Price Changes

Contemporaneous (One-Week Change)	Two-Week Trend	Three-Week Trend	Four-Week Trend	Five-Week Trend
0.2876	0.4500	0.5440	0.5822	0.6202

Table 2
Average Weekly Price Changes in Retail Gasoline and Crude Oil
Cents per Gallon

	Jan. ’91 to Dec. ‘96	Jan. ’97 to May ‘02	Jan. ’97 to Dec. ‘98	Jan. ’99 to May ‘02
Gasoline:				
ΔP	0.01	0.04	-0.30	0.25
Std. Dev.	1.01	2.48	0.88	3.03
Crude Oil:				
ΔP		-0.04	-0.33	0.13
Std. Dev.		2.32	1.45	2.69

Table 3
Augmented Dickey-Fuller Tests for Unit Roots - - Monthly Data

Variable	Type	Test Statistic†	Lags
Regular Gas Price	Levels	2.828***	3
Regular Gas Price	first differences	8.053*	2
Crude Price	Levels	2.313	4
Crude Price	first differences	9.267*	3
Monthly Income	Levels	4.120*	4
Monthly Income	first differences	12.521*	2

*, and *** indicate the null hypothesis of a unit root can be rejected at the 1%, and 10% level of significance, respectively. Lags indicate number of lagged first differences included as explanatory variables in the Augmented Dickey-Fuller (ADF) regression. All specifications include a constant, but no trend. †1% critical value = 3.454, 5% critical value = 2.871, 10% critical value = 2.572. Sample consists of 307 monthly observations between July 1976 and December 2001.

Table 4
The Effect of Changes in the Price of Crude Oil on the Price of Regular Gas: Monthly U.S. Data

	Dependent Variable and Sample			
Independent Variables	**Gas Price$_t$** *(76:2-01:12)*	**Gas Price$_t$** *(76:3-01:12)*	**Change in Gas Price$_t$** *(76:3-01:12)*	**Change in Gas Price$_t$** *(76:4-01:12)*
Intercept	22.866*	31.074*	-1.014	-1.104***
	(1.461)	(3.853)	(0.648)	(0.655)
Crude Price$_t$	2.753*	2.416*		
	(0.042)	(0.114)		
D$_t$ (Crude Price$_t$)	-0.086*	-0.062*		
	(0.027)	(0.015)		
Income$_t$	7.538*	7.059*		
	(0.156)	(0.619)		
Change in Crude Price$_t$			1.922*	1.900*
			(0.228)	(0.251)
D$_t$ (Change in Crude Price$_t$)			0.137	0.151
			(0.361)	(0.390)
Change in Income$_t$			0.506	1.030
			(5.794)	(5.509)
February	0.215	0.228	1.036	1.090
	(1.381)	(0.605)	(0.881)	(0.802)
March	0.830	0.752	1.418	1.515***
	(1.382)	(0.802)	(0.873)	(0.871)
April	2.955**	2.979*	3.402*	3.485*
	(1.380)	(0.919)	(0.869)	(0.870)
May	5.234*	5.265*	3.259*	3.335*
	(1.388)	(0.997)	(0.881)	(0.884)
June	5.893*	6.081*	2.234**	2.310*
	(1.380)	(1.034)	(0.870)	(0.873)
July	4.723*	4.935*	-0.001	0.071
	(1.380)	(1.048)	(0.878)	(0.881)
August	3.604*	3.925*	0.162	0.241
	(1.380)	(1.035)	(0.872)	(0.878)
September	3.183**	3.647*	0.879	0.964
	(1.381)	(0.995)	(0.872)	(0.876)
October	1.693	2.194**	-0.378	-0.300
	(1.382)	(0.922)	(0.872)	(0.874)
November	1.330	1.798**	0.686	0.758
	(1.380)	(0.800)	(0.874)	(0.865)
December	0.553	0.943	0.248	0.304
	(1.380)	(0.603)	(0.883)	(0.809)
ñ		0.849*		0.177*
		(0.032)		(0.062)
R^2	0.956	0.984	0.471	0.487
F	461.660*	1219.137*	18.774*	18.542*
DW	0.404	1.535	1.620	1.855
N	311	310	310	309

Standard errors in parentheses. *, **, and *** indicate the estimated coefficient is statistically different from zero at the 1%, 5%, or 10% level of significance, respectively. All t-tests are two-tail tests. D$_t$ equals 1 if the price of crude increased since the last period, 0 otherwise. The ñ is the first-order autocorrelation coefficient in a AR(1) model.

Table 5
Augmented Dickey-Fuller Tests for Unit Roots - - Weekly Data

Variable	Type	Test Statistic
U.S. Regular Gas Price	levels	2.121
U.S. Regular Gas Price	first differences	6.70*
Crude Price	levels	1.573
Crude Price	first differences	9.755*
Regular Gas Price– East Region	levels	1.929
Regular Gas Price– East Region†	first differences	6.298*
Regular Gas Price – New England Region	levels	2.001
Regular Gas Price – New England Region	first differences	5.107*
Regular Gas Price– Central Atlantic Region	levels	1.974
Regular Gas Price– Central Atlantic Region	first differences	5.121*
Regular Gas Price– Lower Atlantic Region	levels	1.959
Regular Gas Price– Lower Atlantic Region	first differences	6.713*
Regular Gas Price– Midwest Region†	levels	2.892**
Regular Gas Price– Midwest Region	first differences	7.628*
Regular Gas Price– Gulf Region	levels	1.915
Regular Gas Price– Gulf Region	first differences	6.558*
Regular Gas Price– Rockies Region†	levels	2.148
Regular Gas Price– Rockies Region	first differences	6.784*
Regular Gas Price– West Region†	levels	2.380
Regular Gas Price– West Region	first differences	6.782*
Regular Gas Price– California	levels	1.975
Regular Gas Price– California	first differences	3.804*
Regular Gas Price– Colorado	levels	2.322
Regular Gas Price– Colorado	first differences	4.375*
Regular Gas Price– Minnesota	levels	2.855***
Regular Gas Price– Minnesota	first differences	5.302*
Regular Gas Price– New York State	levels	2.739***
Regular Gas Price– New York State	first differences	2.982**
Regular Gas Price– Texas	levels	2.127
Regular Gas Price– Texas	first differences	3.877*
Regular Gas Price– Chicago	levels	4.237*
Regular Gas Price– Chicago	first differences	4.096*
Regular Gas Price– Denver	levels	2.244
Regular Gas Price– Denver	first differences	4.430*
Regular Gas Price– Houston	levels	1.973
Regular Gas Price– Houston	first differences	3.674*
Regular Gas Price– Los Angeles	levels	1.869
Regular Gas Price– Los Angeles	first differences	3.937*
Regular Gas Price– New York City	levels	2.545
Regular Gas Price– New York City	first differences	2.922**
Regular Gas Price– San Francisco	levels	1.764
Regular Gas Price– San Francisco	first differences	3.468**

*,**,*** indicate the null hypothesis of a unit root can be rejected at the 1%, 5%, and 10% level of significance, respectively. † indicates the ADF regression contained two lags, while all others only contained one lag. All specifications include a constant, but no trend. The U.S. and regional level data samples consist of 271 weekly observations between 1/27/97 and 4/1/02. The state and city level data consists of 94 observations between 6/19/-- and 4/1/02.

Table 6
National and Regional Effects of Crude Oil Price on Regular Gas Price: Weekly U.S. Data (Levels)

$$\text{Gas Price}_t = \beta_0 + \beta_1(\text{Crude Price}_t) + \beta_2 D_t(\text{Crude Price}_t) + u_t$$

Gas Prices in:	β_0	β_1	β_2	ρ	R^2	DW
U.S.	64.337*	3.000*	-0.125*		0.87	0.12
	(1.467)	(0.075)	(0.041)			
U.S.	90.579*	1.726*	-0.028*	0.980*	0.99	1.13
	(6.881)	(0.146)	(0.012)	(0.012)		
East	63.051*	3.071*	-0.182*		0.87	0.14
	(1.432)	(0.073)	(0.040)			
East	102.399*	1.229*	-0.027*	0.993*	0.99	0.84
	(14.247)	(0.110)	(0.007)	(0.007)		
New England	67.824*	3.220*	-0.236*		0.86	0.17
	(1.602)	(0.082)	(0.045)			
New England	112.409*	0.958*	-0.028*	0.995*	0.99	0.64
	(18.733)	(0.113)	(0.007)	(0.006)		
Central Atlantic	66.502*	3.170*	-0.209*		0.86	0.15
	(1.563)	(0.008)	(0.044)			
Central Atlantic	110.339*	1.023*	-0.0245*	0.995*	0.99	0.67
	(17.845)	(0.110)	(0.007)	(0.006)		
Lower Atlantic	60.291*	2.967*	-0.154*		0.88	0.13
	(1.380)	(0.071)	(0.039)			
Lower Atlantic	92.945*	1.428*	-0.029*	0.988*	0.99	1.04
	(9.734)	(0.125)	(0.008)	(0.009)		
Midwest	61.150*	3.143*	-0.048		0.79	0.13
	(2.033)	(0.104)	(0.057)			
Midwest	73.597*	2.551*	-0.031***	0.943*	0.97	1.29
	(5.958)	(0.237)	(0.017)	(0.020)		
Gulf	62.993*	2.879*	-0.154*		0.86	0.14
	(1.408)	(0.072)	(0.040)			
Gulf	95.201*	1.320*	-0.024*	0.988*	0.99	1.05
	(9.597)	(0.126)	(0.008)	(0.009)		
Rockies	73.982*	2.894*	-0.219*		0.82	0.18
	(1.659)	(0.085)	(0.047)			
Rockies	111.873*	1.019*	-0.035*	0.984*	0.98	1.31
	(9.851)	(0.169)	(0.011)	(0.011)		
West	80.986*	3.049*	-0.187*		0.74	0.10
	(2.203)	(0.113)	(0.062)			
West	124.496*	1.88*	-0.038*	0.984*	0.99	0.67
	(0.1125)	(0.0018)	(0.0001)	(0.0099)		

Standard errors in parentheses. * and *** indicate the estimated coefficient is statistically different from zero at the 1% or 10% level of significance, respectively. All t-tests are two-tail tests. D_t equals 1 if the price of crude increased since the last period, 0 otherwise. The sample size of the regressions without the ñ associated with the AR(1) error term consists of 273 weekly observation between 1/13/97 and 4/1/02. The regressions that include the AR(1) error specification have 272 observation from 1/20/97 to 4/1/02.

Table 7
National and Regional Effects of Crude Oil Price on Regular Gas Price:
Weekly U.S. Data (First Differences)

$$\Delta(\text{Gas Price})_t = \beta_0 + \beta_1 \Delta(\text{Crude Price})_t + \beta_2 D_t \Delta(\text{Crude Price})_t + e_t$$

Ä Gas Prices in:	β_0	β_1	β_2	ρ	R^2	DW
U.S.	-0.225	1.186*	0.744***		0.34	1.09
	(0.184)	(0.211)	(0.387)			
U.S.	-0.215	0.611*	0.732**	0.570*	0.52	2.19
	(0.269)	(0.175)	(0.303)	(0.051)		
East	-0.244***	0.697*	0.777*		0.30	0.79
	(0.140)	(0.168)	(0.295)			
East	-0.151	0.233**	0.589*	0.738*	0.64	2.03
	(0.268)	(0.105)	(0.181)	(0.042)		
New England	-0.186	0.523*	0.552***		0.18	0.60
	(0.144)	(0.165)	(0.303)			
New England	-0.050	0.175***	0.283***	0.789*	0.66	2.08
	(0.303)	(0.095)	(0.163)	(0.038)		
Central Atlantic	-0.212	0.569*	0.646**		0.23	0.61
	(0.140)	(0.160)	(0.294)			
Central Atlantic	-0.090	0.189**	0.437*	0.791*	0.68	2.06
	(0.298)	(0.092)	(0.159)	(0.038)		
Lower Atlantic	-0.271***	0.836*	0.877*		0.32	1.00
	(0.158)	(0.182)	(0.333)			
Lower Atlantic	-0.200	0.334**	0.711*	0.637*	0.55	2.09
	(0.254)	(0.139)	(0.242)	(0.048)		
Midwest	-0.247	1.875*	0.833		0.28	1.29
	(0.314)	(0.360)	(0.662)			
Midwest	-0.290	1.259*	0.948	0.408*	0.38	2.12
	(0.392)	(0.340)	(0.592)	(0.057)		
Gulf	-0.313**	0.723*	0.971*		0.29	1.00
	(0.158)	(0.181)	(0.333)			
Gulf	-0.286	0.245***	0.901*	0.6119*	0.53	2.15
	(0.242)	(0.142)	(0.247)	(0.049)		
Rockies	-0.123	0.559**	0.453		0.09	1.25
	(0.217)	(0.248)	(0.456)			
Rockies	-0.028	0.239	0.240	0.432*	0.23	2.14
	(0.275)	(0.228)	(0.401)	(0.432)		
West	-0.086	0.574**	0.507		0.08	0.65
	(0.235)	(0.270)	(0.495)			
West	0.132	0.295***	-0.081	0.729*	0.55	1.82
	(0.425)	(0.172)	(0.298)	(0.042)		

Standard errors in parentheses. *, **, and *** indicate the estimated coefficient is statistically different from zero at the 1%, 5%, or 10% level of significance, respectively. All t-tests are two-tail tests. D_t equals 1 if the price of crude increased since the last period, 0 otherwise. The sample size of the regressions without the ñ associated with the AR(1) error term consists of 272 weekly observation between 1/20/97 and 4/1/02. The regressions that include the AR(1) error specification have 271 observation from 1/27/97 to 4/1/02.

Table 8

The Effect of Changes in the Crude Price on Gas Prices in Selected States and Cities: Weekly Data, Not Corrected for Autocorrelation

Gas Prices in:	***Using level data:*** $\text{Gas Price}_t = \beta_0 + \beta_1(\text{Crude Price}_t) + \beta_2 D_t(\text{Crude Price}_t) + v_t$					***Using first-differences of data:*** $\Delta(\text{Gas Price})_t = \beta_0 + \beta_1\Delta(\text{Crude Price})_t + \beta_2 D_t\Delta(\text{Crude Price})_t + u_t$				
States	β_0	β_1	β_2	R^2	DW	β_0	β_1	β_2	R^2	DW
California	72.404*	3.787*	-0.277**	0.59	0.14	-0.144	0.901**	0.358	0.12	0.74
	(7.887)	(0.326)	(0.115)			(0.550)	(0.442)	(0.919)		
Colorado	61.266*	3.552*	-0.290*	0.63	0.19	-0.330	0.979**	0.461	0.13	0.88
	(6.818)	(0.282)	(0.099)			(0.604)	(0.486)	(1.009)		
Minnesota	62.469*	3.333*	-0.002	0.51	0.21	-0.551	2.211**	1.083	0.19	1.55
	(8.408)	(0.348)	(0.122)			(1.074)	(0.864)	(1.795)		
New York	74.539*	3.258*	-0.237*	0.66	0.16	-0.525	0.419	0.782	0.15	0.41
	(5.835)	(0.241)	(0.085)			(0.351)	(0.282)	(0.587)		
Texas	58.291*	3.110*	-0.143***	0.68	0.11	-0.794	0.759**	1.464**	0.32	0.99
	(5.466)	(0.226)	(0.079)			(0.408)	(0.328)	(0.682)		
Cities										
Chicago	71.370*	3.226*	-0.091	0.34	0.13	-1.204	1.669***	1.742	0.14	1.46
	(11.548)	(0.478)	(0.168)			(1.165)	(0.937)	(1.947)		
Denver	63.851*	3.415*	-0.314*	0.58	0.22	-0.393	0.728	0.620	0.08	1.17
	(7.296)	(0.302)	(0.106)			(0.674)	(0.542)	(1.127)		
Houston	54.320*	3.292*	-0.165***	0.65	0.11	-0.735	0.468	1.296***	0.17	1.03
	(6.168)	(0.255)	(0.090)			(0.446)	(0.359)	(0.746)		
Los Angeles	67.819*	3.630*	-0.240***	0.48	0.10	-0.292	0.650	0.789	0.08	0.86
	(9.563)	(0.396)	(0.139)			(0.647)	(0.521)	(1.081)		
New York City	68.062*	3.473*	-0.257*	0.65	0.14	-0.486	0.467	0.518	0.14	0.48
	(6.463)	(0.267)	(0.094)			(0.353)	(0.284)	(0.589)		
San Francisco	85.062*	3.835*	-0.338*	0.58	0.18	-0.313	0.670	0.468	0.09	0.79
	(8.130)	(0.336)	(0.118)			(0.536)	(0.432)	(0.897)		

Standard errors in parentheses. *, **, and *** indicate the estimated coefficient is statistically different from zero at the 1%, 5%, or 10% level of significance, respectively. All t-tests are two-tail tests. D_t equals 1 if the price of crude increased since the last period, 0 otherwise. The sample size of the regressions with the level data consists of 96 observations between 6/5/00 and 4/1/02. The regressions with differenced data consist of 95 observations between 6/12/00 and 4/1/02.

Table 9

The Effect of Changes in the Crude Price on Gas Prices in Selected States and Cities: Weekly Data, Corrected for Autocorrelation

	Using Level Data:						***Using First-Differenced Data***					
	$\text{Gas Price}_t = \beta_0 + \beta_1(\text{Crude Price}_t) + \beta_2 D_t(\text{Crude Price}_t) + v_t$						$\Delta(\text{Gas Price})_t = \beta_0 + \beta_1\Delta(\text{Crude Price})_t + \beta_2 D_t \Delta(\text{Crude Price})_t + u_t$					
	β_0	β_1	β_2	ñ	R^2	DW	β_0	β_1	β_2	ñ	R^2	DW
States												
California	132.418*	1.281*	-0.031	0.981*	0.98	0.78	0.091	0.500***	-0.074	0.689*	0.52	1.95
	(20.167)	(0.335)	(0.023)	(0.020)			(0.909)	(0.290)	(0.555)	(0.077)		
Colorado	103.374*	1.575*	-0.054**	0.971*	0.97	0.96	-0.335	0.339	0.175	0.641*	0.46	1.89
	(16.033)	(0.363)	(0.025)	(0.026)			(0.924)	(0.342)	(0.657)	(0.080)		
Minnesota	69.530*	3.057*	-0.054	0.898*	0.91	1.46	-0.729	1.886**	0.877	0.207**	0.24	2.00
	(16.312)	(0.626)	(0.045)	(0.047)			(1.130)	(0.836)	(1.648)	(0.103)		
New York	112.436*	0.924*	-0.029***	0.990*	0.98	0.46	-0.394	0.181	0.308	0.852*	0.76	1.74
	(37.366)	(0.214)	(0.015)	(0.016)			(0.831)	(0.121)	(0.231)	(0.055)		
Texas	89.227*	1.612*	-0.037**	0.974*	0.98	0.98	-0.636	0.300	1.081**	0.647*	0.57	2.12
	(12.236)	(0.250)	(0.017)	(0.023)			(0.648)	(0.236)	(0.453)	(0.081)		
Cities												
Chicago	76.045*	2.824*	-0.076	0.922*	0.91	1.38	-1.286	1.105	1.321	0.300*	0.22	2.12
	(18.756)	(0.685)	(0.049)	(0.037)			(1.309)	(0.885)	(1.746)	(0.101)		
Denver	109.822*	1.286*	-0.040	0.967*	0.96	1.26	-0.447	0.257	0.501	0.455*	0.25	2.06
	(16.577)	(0.411)	(0.028)	(0.028)			(0.846)	(0.467)	(0.0091)	(0.093)		
Houston	96.676*	1.259*	-0.041**	0.978*	0.98	1.00	-0.625	-0.002	0.874***	0.604*	0.44	2.14
	(15.340)	(0.271)	(0.019)	(0.021)			(0.658)	(0.269)	(0.517)	(0.084)		
Los Angeles	128.106*	1.220*	-0.035	0.979*	0.97	0.88	-0.030	0.291	0.255	0.620*	0.42	2.02
	(22.269)	(0.395)	(0.027)	(0.021)			(0.962)	(0.377)	(0.726)	(0.083)		
New York City	95.518	0.962*	-0.038**	0.993*	0.99	0.47	-0.376	0.156	0.111	0.839*	0.73	1.88
	(75.414)	(0.211)	(0.014)	(0.014)			(0.804)	(0.128)	(0.244)	(0.056)		
San Francisco	144.778*	0.955*	-0.014	0.986*	0.98	0.80	-0.029	0.383	-0.006	0.654	0.45	2.06
	(29.071)	(0.333)	(0.023)	(0.019)			(0.845)	(0.300)	(0.576)	(0.082)		

Standard errors in parentheses. *, **, and *** indicate the estimated coefficient is statistically different from zero at the 1%, 5%, or 10% level of significance, respectively. All t-tests are two-tail tests. D_t equals 1 if the price of crude increased since the last period, 0 otherwise. The coefficient associated with the AR(1) error specification is reported under the column with a heading of ñ. The sample size of the regressions with the level data consists of 95 observations between 6/12/00 and 4/1/02. The regressions with differenced data consist of 94 observations between 6/19/00 and 4/1/02.

Table 10
Augmented Dickey-Fuller Tests for Unit Roots - - Daily Spot Prices

Variable	Type	Test Statistic
Daily Spot Regular Gas Price	levels	3.971*
Daily Spot Crude Price	levels	3.798*

* indicates the null hypothesis of unit root is rejected at the 1% level. Sample consists of 3986 weekday observations between 65/86 and 4/1/02.

Table 11
The Effect of Changes in the Price of Crude Oil on the Price of Regular Gas: Using U.S. Daily Spot Prices

Using Level Data:

$$\text{Regular Gas Price}_t = \beta_0 + \beta_1(\text{Crude Price}_t) + \beta_2 D_t(\text{Crude Price}_t) + u_t$$

β_0	β_1	β_2	ρ	R^2	DW
4.035*	2.692*	-0.001		0.859	0.058
(0.361)	(0.018)	(0.008)			
19.307*	1.943*	-0.001	0.982*	0.993	1.806
(1.315)	(0.037)	(0.001)	(0.003)		

Using First Differenced Data:

$$\Delta(\text{Regular Gas Price})_t = \beta_0 + \beta_1\Delta(\text{Crude Price}_t) + \beta_2 D_t\Delta(\text{Crude Price})_t + e_t$$

α_0	α_1	α_2	ρ	R^2	DW
-0.031	1.829*	0.185**		0.455	1.820
(0.024)	(0.050)	(0.086)			
-0.031	1.812*	0.192**	0.090*	0.460	2.00
(0.026)	(0.050)	(0.087)	(0.016)		

Standard errors in parentheses. * and ** indicate the estimated coefficient is statistically different from zero at the 1%, or 5% level of significance, respectively. All t-tests are two-tail tests. D_t equals 1 if the price of crude increased since the last period, 0 otherwise. The coefficient associated with the AR(1) error specification is reported under the column with a heading of ñ. The sample of spot prices consists of weekday observations between 6/4/86 and 4/1/02.

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Appendix 1

Regions that compose the East Coast region

The New England region The Central Atlantic region The Lower Atlantic region

States in the New England Region:

Connecticut New Hampshire
Maine Rhode Island
Massachusetts Vermont

States in the Gulf Region:

Alabama Mississippi
Arkansas New Mexico
Louisiana Texas

States in the Central Atlantic Region:

Delaware New York
Maryland Pennsylvania
New Jersey

States in the Rockies Region:

Colorado Utah
Idaho Wyoming
Montana

States in the Lower Atlantic Region:

Florida South Carolina
Georgia Virginia
North Carolina West Virginia

States in the West Region:

Alaska Nevada
Arizona Oregon
California Washington
Hawaii

States in the Midwest Atlantic Region:

Illinois Nebraska
Indiana North Dakota
Iowa Ohio
Kansas Oklahoma
Kentucky South Dakota
Michigan Tennessee
Minnesota Wisconsin
Missouri