DOES OPEC MATTER?
AN ECONOMETRIC ANALYSIS OF OIL PRICES

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ABSTRACT

In this paper, we assess claims that Opec's ability to influence real oil prices has diminished and that the negative relation between real oil prices and Opec production represents a backward bending supply curve. To do so, we estimate an equation for real oil prices from quarterly data, 1986:III through 2000:III. The results indicate that there is a statistically significant relation among real oil prices, Opec capacity utilization, Opec quotas, the degree to which Opec exceeds these production quotas, and OECD stocks of crude oil. Further analysis indicates that these variables 'Granger cause' real oil prices but real oil prices do not 'Granger cause' Opec capacity utilization, Opec production quotas, the degree to which they exceed those production quotas, and oil stocks in OECD nations. These results indicate that Opec plays an important role in determining real oil prices. The analysis of causal order indicates that the negative relation between price and production is part of the cointegrating relation for oil prices, not oil production. This implies that the previous models cannot be used to test competing models for Opec production behavior. The effect of OECD oil stocks on real oil prices indicates that the private savings associated with recent reductions in inventories may be less than the social costs associated with higher oil prices. As such, there may be an important externality in private decisions regarding optimal crude oil stocks.

Introduction

The role of Opec in world oil markets has been examined by both the press and the academic community over the last few years. Recent stories in the press speculate that Opec's ability to affect real prices has diminished (e.g. Belton, 1998). They point out that the general decline in real oil prices during much of the 1990's coincides with an increase in Opec's market share. In addition, the press describes a tendency for Opec members to exceed their production quotas.

In addition to these topics, the academic community is interested in the nature and goals of Opec. When examining Opec production, Griffin (1985) and subsequent analysts test four competing hypotheses for Opec behavior; (1) a competitive model, (2) a cartel model, (3) a target revenue model, and (4) a property rights model. As described below, efforts to choose among these behaviors focus, in part, on the relation between price and Opec production.

In order to investigate the above concerns about Opec, we estimate an equation for real oil prices from quarterly data, 1986:III through 2000:III. The choice of the sample period is determined by economic considerations and data constraints. The beginning of the sample period is defined by the "end" of the 1986 price collapse, when Opec reinstituted quotas after changing its announced strategy from defending an official price to defending market share. The closing date of the sample period is given by the absence of comprehensive data on Opec capacity after 2000:III. The results indicate that the stochastic trends in real oil prices cointegrate with stochastic trends in Opec capacity utilization, Opec production quotas, the degree to which Opec exceeds these quotas, and crude oil stocks in OECD nations. Further analysis indicates that these variables

"Granger cause" real oil prices but real oil prices do not "Granger cause" Opec capacity utilization, Opec production quotas, the degree to which they cheat on those quotas, or oil stocks in OECD nations.

These results contrast with claims about the decline in Opec market power and the way in which analysts interpret the negative relation between oil prices and Opec production. The cointegrating relation among real oil prices, Opec capacity utilization, Opec quotas, and adherence to those quotas indicates that Opec plays an important role in determining world oil prices. Our analysis of causal order for individual Opec countries shows that there is much more convincing evidence that production 'Granger causes' prices than the other way around. This result indicates that the negative relation between prices and production is best interpreted to be part of a cointegrating relation for price, not production. Finally, the role of OECD oil stocks indicates that private decisions to reduce inventories may ignore an important externality that reduces private costs but raises social costs.

These results and the methods used to obtain them are described in five sections. In the second section, we describe the data and econometric methods that are used to estimate the relation and causal order among real oil prices, Opec behavior, and OECD stocks. These results are described in section III. In section IV, we discuss the results relative to Opec's role in setting real oil prices, Opec production behavior, and private decisions about optimal stocks. The limits associated with this research are described in the conclusion (section V).

II Methods

In this section, we describe three estimators that are used to investigate the relation among real oil prices, Opec behavior, and OECD stocks. Next, we describe two methods that are used to examine causal relations among these variables. Finally, we describe models that are used to test competing hypotheses about production behavior by Opec and non-Opec nations.

A cointegrating relation for real oil prices

We estimate a model for real oil prices that includes variables which represent market conditions, such as OECD stocks of crude oil, capacity utilization by Opec, and non-competitive behaviors by Opec such as Opec announced quotas and the degree to which its production adheres to these quotas. These variables are specified using a model that builds on a previous effort by Kaufmann (1995) as follows:

$$\text{Price} = \alpha + \beta_1 * \text{Days} + \beta_2 * \text{Quota} + \beta_3 * \text{Cheat} + \beta_4 * \text{Caputil} +$$

$$\beta_5\, \text{Q1} + \beta_6 * \text{Q2} + \beta_7 * \text{Q3} + \beta_8 * \text{War} + \mu \qquad (1)$$

in which:

Price is the US crude oil import FOB price and is measured in 1996 US $ per barrel, *Days* is days of forward consumption of OECD crude oil stocks, which is calculated by dividing OECD crude oil stocks by OECD crude oil demand, *Quota* is the Opec production quota (million barrels per day), *Cheat* is the difference between Opec crude oil production and Opec quotas (million barrels per day), *Caputil* is capacity utilization by Opec, which is calculated by dividing Opec production (mbd) by Opec capacity (mbd), *Q1*, *Q2*, and *Q3* are dummy variables for quarters I, II, and III, respectively, and *War* is a dummy variable for the Persian Gulf War (third and fourth quarters of 1990). Opec data (production, capacity, and Quota) include all member nations at time t (e.g.

Ecuador is included through the fourth quarter of 1992, when it formally withdrew from Opec and Gabon up to its formal withdrawal in the fourth quarter of 1994). The regression results described in section III do not change significantly if Opec is re-defined to include its current eleven members throughout the sample period.

Estimation technique

To test variables for unit roots and to test for cointegration among variables, we use the augmented Dickey Fuller (ADF) statistic (Dickey and Fuller, 1979) and the procedure developed by Hylleberg *et al*., (1990) (HEGY). The results in Table 1 indicate that Quota, Cheat, and Caputil contain an annual root (we fail to reject both the ADF and $\pi_1 = 0$) but do not contain a seasonal root ($\pi_2 = 0$, $\pi_3 = \pi_4 = 0$). On the other hand, HEGY tests of Price and Days reject the null hypothesis ($\pi_1 = 0$) at the five percent level, but not the one percent level. If these results are correct, they imply that the cointegrating relation described in the next section includes Quota, Cheat, and Caputil only. We evaluate this notion by estimating a version of equation (1) in which Price is replaced by one of the three I(1) variables and Days is eliminated from the right hand side. Regardless of the dependent variable, the residuals from all three regression fail to reject $\pi_1 = 0$ at even the ten percent level. This result indicates that Quota, Cheat, and Caputil do not cointegrate. Based on results described in section III, which indicate that the five variables in equation (1) constitute a cointegrating relation, this result indicates that Price and Days also are I(1).

The presence of I(1) trends invalidates the blind application of ordinary least squares (OLS) because the diagnostic statistics generated by OLS indicate a meaningful relation among unrelated I(1) variables more often than implied by random chance (Granger and

Newbold, 1974). Such relations are termed spurious regressions. To avoid the confusion that is associated with spurious regressions, the relation among variables in equation (1) is evaluated by determining whether they cointegrate. Following a well-established method to determine whether two (or more) variables cointegrate (Engle and Granger, 1987), OLS is used to estimate equation (1) and the residual (μ) is analyzed for a stochastic trend using the ADF and HEGY statistics. If the ADF or HEGY statistic fails to reject the null hypothesis, the nonstationary residual indicates that the regression is spurious. If the regression residual is stationary, the variables cointegrate. In this case, there is a statistically meaningful long-run relation among the I(1) variables.

Even if the variables cointegrate, the OLS estimate of the cointegrating vector will contain a small-sample bias and the limiting distribution will be non-normal with a non-zero mean (Stock, 1987). To avoid confusion associated with this bias, we use statistical techniques that are designed to analyze the relation among nonstationary variables. These techniques include the dynamic ordinary least squares (DOLS) estimator developed by Stock and Watson (1993) and the full information maximum likelihood (FIML) estimator of a vector error correction model developed by Johansen (1988) and Johansen and Juselius (1990). Rather than rely on a single estimation technique, we evaluate the degree to which the results are robust by using both efficient estimators and OLS to estimate the relation between Price and the explanatory variables in equation (1). If these variables cointegrate, the point estimates for the long-run relation should be similar across the three estimation techniques. If there is disagreement, results generated by techniques designed for the analysis of integrated variables (DOLS, FIML) are preferred.

We use DOLS to estimate the cointegrating relation given by equation (1) because it generates asymptotically efficient estimates of the regression coefficients for variables that cointegrate, it is computationally simple, and it performs well relative to other asymptotically efficient estimators (Stock and Watson, 1993). The coefficients estimated by DOLS represent the long-run relation among variables. DOLS does not estimate the short-run dynamics -- it is not necessary for asymptotically efficient estimation of the cointegrating relation. The lags and leads used to estimate the DOLS version of equation (1) are chosen using the Akaike Information criterion (Akaike, 1973).

To examine the short run relation among variables in equation (1), we use OLS to estimate an error correction model (ECM), which is given by equation (2):

$$\Delta Price_t = k + \alpha \mu_{t-1} + \sum_{i=1}^{s} \lambda_{1i} \Delta Days_{t-i} + \sum_{i=1}^{s} \lambda_{2i} \Delta Quota_{t-i} + \sum_{i=1}^{s} \lambda_{3i} \Delta Cheat_{t-i} + \sum_{i=1}^{s} \lambda_{4i} \Delta Caputil_{t-i} + \sum_{i=1}^{s} \lambda_{5i} \Delta Price_{t-i} + \varepsilon_t \qquad (2)$$

in which Δ is the first difference operator (e.g. $Price_t - Price_{t-1}$) and μ is the regression residual from equation (1) that is estimated using either OLS or DOLS. The statistical significance of α indicates if disequilibrium between real oil prices and the right hand side variables in equation (1) affects price. Disequilibrium is given by μ, which represents the difference between the observed value for real oil prices and the long-run value that is implied by the variables on the right-hand side equation (1). If this disequilibrium affects price, there will be a statistically significant relation between μ and the subsequent change in price. Under these conditions, the right-hand side variables in equation (1) are said to "Granger cause" real oil prices. The number of lags (s) for the right-hand side variables in equation (2) is chosen using the Akaike information criterion

(Akaike, 1973).

To estimate the long and short run relation between real oil prices and the explanatory variables simultaneously, we use a full information maximum likelihood procedure to estimate a partial system for a vector error correction model (VECM) in which there is only one equation that is specified as follows:

$$\Delta \text{Price}_t = \alpha(\beta_1 \text{Price}_{t-1} + \beta_2 X_{t-1}) + \sum_{i=1}^{s} \lambda_i \Delta \text{Price}_{t-i} + \sum_{i=0}^{s} \varphi_i \Delta X_{t-i} + \sum_{i=0}^{s} \theta_i Z_{t-i} + \varepsilon_t \quad (3)$$

in which X is a vector of the right-hand side variables from equation (1) and Z is the vector of dummy variables (Q1, Q2, Q3, and War). Equation (3) specifies the first difference of real oil prices, which is stationary, as a linear function of linear lagged values for the first difference of the nonstationary variables, which also are stationary, and stationary combinations of the nonstationary variables, which represent the long term relations(s) among variables. The long run relation among variables is given by the elements of β (β_1, β_2) and is termed the cointegrating vector. The rate at which real oil prices respond to disequilibrium in the long run relation is given by α. The interpretation of α in equation (3) is similar to that of α from equation (2).

Causal order

We also wish to determine whether the relation among variables in equation (1) is coincidental or whether there is a “dependent” variable that is meaningfully dependent on changes in the “independent” variables. This type of dependence can be examined by testing for “Granger causality” (Granger, 1969). To evaluate the direction of “Granger causality” among real oil prices and the right hand side variables from equation (1), we estimate a VAR with five endogenous variables; Prices, Days, Caputil, Cheat, and Quota and four I(0) dummy variables; Q1, Q2, Q3, and War. The lag length of the VAR is chosen using the Sims (1980) Likelihood ratio test. The direction of causal order is determined by restricting the lagged values of an endogenous variable to zero. Rejecting the null hypothesis that the coefficients are zero (i.e. the excluded variable does not Granger cause the dependent variable) indicates that the lagged values of the excluded variable have information about the dependent variable beyond that contained in lagged values of the dependent variable and the other right hand side variables. This would indicate that the excluded variable "Granger causes" the left-hand-side variable. To ensure completeness, we test restrictions that evaluate all possible directions of Granger causality.

Restrictions on the lagged value(s) of a variable are evaluated with a F test. This statistic can be evaluated reliably against an F distribution if the variables in the VAR cointegrate—if they do not cointegrate, the test statistic will reject the restriction more often than expected by random chance (Ohanian, 1988; Toda and Phillips, 1993). Under these conditions, the restrictions are more likely to indicate a direction of causality when none may be present.

The degree to which the direction of Granger causality indicated by the VAR is robust is evaluated by imposing restrictions on a VECM. The VECM specifies the same five endogenous variables; Prices, Days, Caputil, Cheat, and Quota and the same four I(0) dummy variables; Q1, Q2, Q3, and War. After determining the lag length (using Hannon Quinn statistic) and the number of cointegrating relations (using the λ_{max} and λ_{trace} statistics), the direction of causality is evaluated by restricting the elements of α for an equation equal to zero. The failure to reject these restrictions would indicate that disequilibrium in any of the cointegrating relations does not affect that variable. Under these conditions, the variable is said to be exogenous. This would imply that the exogenous variable is not "Granger caused" by any cointegrating combination of variables. Conversely, rejecting the restrictions on α would indicate that disequilibrium in one or more of the cointegrating relations affects the equation for that variable. This would imply that the variables in the cointegrating relation that affects the equation for a particular variable "Granger cause" that variable. For example, rejecting $\alpha = 0$ in equation (3) (equation 3 would be the equation for price in the full VECM) would indicate that disequilibrium in the cointegrating relation affects real oil prices, which would indicate that the variables in the cointegrating relation "Granger cause" real oil prices.

Opec behavior at the country level

Analyses in previous papers interpret the relation between price and production to evaluate competing models for Opec behavior. This approach originates with Griffin (1985), who uses data from individual Opec nations to estimate the following equation:

$$\ln Q_{it} = \alpha_i + \gamma_i \ln P_t + \beta_i \ln Q_{it}^{OO} + \varepsilon_{it} \qquad (4)$$

in which Q is oil production by Opec nation i at time t, P is the real price of oil, and Q^{OO} is oil production by Opec nations other than nation i. Griffin (1985) uses the value of β and γ to test hypotheses about market sharing behavior. Constant market shares imply β=1.0. If Opec allocates market shares among members, but does not use constant shares, $\beta \neq 0$. The value of γ is interpreted as the effect of prices on production.

To evaluate whether β and γ can be used to evaluate competing models of Opec behavior, we investigate the causal relation between price and production by individual

Opec nations. The ADF and HEGY tests of regression residual for equation (4) that are estimated with our sample data and dummy variables to account for seasonal effects indicate that the residual contains an I(1) unit root (i.e. price and production do not cointegrate). To avoid spurious regression results, we modify equation (4) to include the variables that cointegrate with price as follows:

$$\ln Q_{it} = \alpha_i + \gamma_i \ln P_t + \beta_i \ln Q_{it}^{OO} + \lambda \ln K_{it} + \phi \ln U_t^{OO}$$
$$\pi \ln Days_t + \eta_1 Q1 + \eta_2 Q2 + \eta_3 Q3 + \eta_4 War + \varepsilon_{it} \qquad (5)$$

in which K is capacity for Opec nation i at time t, and U is Opec capacity utilization (for Opec nations other than nation i). We use capacity for nation i, as opposed to capacity utilization by nation i, and eliminate Quota to alleviate concerns about simultaneity between production by nation i and capacity utilization by nation i and/or the production quota for nation i. Although we focus on the value of γ, equation (5) includes oil production by other Opec nations because Dahl and Yücel (1991) argue that its inclusion is more efficient than testing hypotheses for competitive and market sharing behavior separately.

Equation (5) is estimated using OLS from data (1986:III-2000:III) for ten Opec nations (Iraq is not included due to its uncertain role in Opec since the end of the Gulf War). To interpret the estimate for γ_i, we examine the direction of causality between price and production in a VAR that includes the six I(1) endogenous variables (Q, P, Q^{OO}, K, U, and D) and the four I(0) dummy variables (Q1, Q2, Q3, and War) in equation (5). The length of the VAR is determined using the Sims likelihood ratio.

To validate conclusions about the price response of Opec production, we analyze the relation between prices and production by non-Opec nations. Again, the lack of cointegration motivates us to expand Griffin's (1985) original specification to include variables that cointegrate with price as follows:

$$\ln Q_{it}^{n} = \alpha + \gamma \ln P_t + \beta_i \ln Q_{it}^{O} + \lambda \ln Days_t + \varphi \ln Quota_t + \pi \ln Cheat_t + \theta \ln Caputil_t + \eta_1 Q1 + \eta_2 Q2 + \eta_3 Q3 + \eta_4 War + \varepsilon_{it} \quad (6)$$

in which Q_{it}^{n} is oil production by non-Opec nation i at time t, and Q^{O} is Opec production at time t, and γ is interpreted as in equation (5). Equation (6) is estimated using OLS for six non-Opec nations; Canada, China, Mexico, Norway, the UK, and the US. Again, we examine the direction of causality between price and production in a VAR that includes the six I(1) endogenous variables (Q, P, Q^{OO}, K, U, and D), and four I(0) dummy variables (Q1, Q2, Q3, and War) from equation (6). The length of the VAR is determined using the Sims likelihood ratio test.

III Results

Regression results for equation (1) indicate that the regression coefficients: (1) constitute a cointegrating vector; (2) are statistically different from zero; and (3) have signs that are consistent with theory. The ADF and HEGY statistics strongly ($p < .01$) reject the null hypothesis that the residual (μ) from equation (1) has a stochastic trend at either the annual or subannual frequencies (Table 1). This result indicates that the variables in equation (1) cointegrate. As described earlier, this conclusion is reinforced by the results from the FIML estimate of equation (3). The λ_{max} statistic (76.5) strongly rejects ($p < .001$) the null hypothesis that equation (3) contains zero cointegrating relations (Osterwald-Lenum, 1992).

There is little evidence of parameter instability. To test the null hypothesis that the regression coefficients associated with the four right hand side I(1) variables are stable, we estimate the Quandt Likelihood Ratio (repeated versions of the Chow test) with 15 percent trimming. The largest value exceeds the $p < .05$ critical value for the first observation beyond the trimming, which is the fourth quarter of 1988 (Figure 1). After this date, there is no evidence for parameter instability. This stability undermines the claim that the explanatory power of capacity utilization has declined over time (Powell, 1990). Consistent with the presence of cointegration, we reject the null hypothesis of no cointegration ($p < .05$) against the alternative hypothesis of cointegration with a one-time shift in the intercept or the slope of the individual I(1) variables using the procedure described by Gregory and Hansen (1996).

The signs on the regression coefficients are the same across the three estimation techniques and are consistent with theory (Table 2). Consistent with results described by Kaufmann (1995) and Balabanoff (1995), the regression coefficient associated with Days is negative. An increase in stocks reduces real oil price by diminishing reliance on current production and thereby reducing the risk premium associated with a supply disruption. Similarly, an increase in the Opec quota tends to alleviate upward pressure on prices. An increase in the Cheat variable also tends to reduce price. An increase in Opec production relative to the quota increases supply relative to the demand perceived by Opec when setting the quota (perceived demand may not be the most important or only variable used to set the quota). Consistent with the notion of target capacity utilization pricing by Opec (e.g. Gately and Kyle, 1977; Kaufmann, 1995), the sign on the regression coefficient associated with Caputil is positive. The positive sign indicates that increases in capacity utilization tend to increase prices. This effect is consistent with hypothesis that Opec was acting as a marginal producer from 1986:III to 2000:III. During this period, Opec tried to set production to match the expected difference between demand and non-Opec supply, which is determined largely by non-Opec capacity (as price takers, non-Opec producers generally operate at or near capacity). As the call for oil from Opec increases production relative to capacity, utilization rates rise, which signals a "tightness" in the market. The War variable has a positive effect on prices—prices rise after the Iraqi invasion of Kuwait in anticipation of a supply disruption, but this effect disappears during the first quarter of 1991, when it is apparent that the war has little effect on oil supplies from the Persian Gulf (Gately, 1995).

Regression results indicate that real oil prices adjust to the variables on the right hand side of equation (1). In all cases, the rate of adjustment (α) is statistically significant. This result is consistent with the interpretation of equation (1) as a cointegrating relation in which the right hand side variables "Granger cause" real oil prices. The value for α (0.56) that is estimated using the DOLS residual is not statistically different from 1.0 ($t = 1.63$, $p < 0.10$). The value for α (0.79) that is estimated from the FIML estimate for equation (3) residual is larger but is statistically different from 1.0 ($t = 3.34$, $p < 0.002$). These large values indicate that prices adjust to their equilibrium value quickly, perhaps instantly, which would be consistent with a forward-looking rational market.

Causal relations among variables

The restrictions on the VAR indicate that the causal order runs from Days, Caputil, and Quota to real oil prices (Table 3). There is no evidence for a causal relation from Cheat to real prices. Nor is there evidence for a causal relationship from real prices to the Cheat variable, or any other right hand side variable from equation (1). Finally, there is no evidence for a causal relationship between any of the other variables. These results are consistent with those described by Gulen (1997).

Tests of restrictions on the VAR generally are consistent with restrictions imposed on the full VECM. The Hannon Quinn statistic indicates that four lags is optimal for the VECM. The λ_{max} and λ_{trace} statistics indicate that the VECM has two cointegrating relations. We fail to reject the null hypothesis that the two elements of α (two cointegrating relations) are equal to zero for three variables; Days, Quotas, and Cheat (Table 4). Under these conditions, disequilibrium in one or both of the two cointegrating relations does not affect the equations for ΔDays, ΔQuota, and ΔCheat. This result indicates that Day, Quota, and Cheat are not “Granger caused” by real oil prices or any of the other I(1) variables in the VECM.

Based on these results, we respecify the VECM with two endogenous variables, Price and Caputil, three exogenous variables, Day, Quota, and Cheat, and four dummy variables, Q1, Q2, Q3, and War. The Hannon Quinn statistic indicates that four lags is optimal for the VECM. The λ_{max} and λ_{trace} statistics indicate that the VECM has two cointegrating relations. This VECM is used to examine the possibility that Caputil “Granger causes” Price, Price “Granger causes” Caputil, and retest the notion that the remaining I(1) variables “Granger Cause” Price.

We reject the null hypothesis that the elements of $\alpha = 0$ for the ΔCaputil equation. The inability to make Caputil exogenous indicates that disequilibrium in at least one of the two cointegrating relations affects the equation for ΔCaputil. To determine whether real oil prices “Granger cause” Opec capacity utilization, we impose a restriction that eliminates Price from one of the two cointegrating relations. The element of α for the

ΔCaputil equation that is associated with the cointegrating relation without price is statistically significant (α = -0.065, t – 3.27) while the element of α that is associated with the cointegrating relation that contains price is not statistically different from zero at $p < .05$ (α = -0.004, t – 1.76). This indicates that real oil prices do not "Granger cause" Opec capacity utilization.

Finally, we reject the null hypothesis that the elements of $\alpha = 0$ for the Δprice equation. The inability to make Price exogenous indicates that disequilibrium in at least one of the two cointegrating relations affects the equation for ΔPrice. To identify which if any of the variables in the VECM affect the Price equation (and support the causal order indicated by the FIML estimate of equation 3), we impose restrictions on the VECM in which one variable is eliminated from one of the cointegrating relations. In each case, both elements of α for the ΔPrice equation are statistically different from zero. Combined with results that indicate no I(1) variable can be eliminated from cointegration space, these results indicate that the VECM is consistent with the hypothesis that Caputil, Quota, Cheat, and Days "Granger cause" real oil prices.

IV Discussion

The results described above indicate that: (1) there is a statistically significant relation among real oil prices, Opec capacity utilization, Opec quotas, the degree to which Opec cheats on these quotas, and OECD stocks; and (2) the direction of causality generally runs from the variables on the right hand side of equation (1) to real oil prices. In this section, we explore these results relative to: (1) Opec's ability to influence oil prices; (2) the relationship between oil prices and Opec production; and (3) an externality in private decisions to set optimal stocks. This discussion indicates that; (1) Opec's ability to influence prices may extend beyond its ability to affect supply; (2) previous efforts to interpret the relation among production by individual Opec nations, other Opec nations, and price may be invalid and therefore these previous conclusions about Opec behaviors are flawed; and (3) the private savings associated with reductions in OECD stocks may be less than the social costs associated with higher oil prices.

Opec affects real oil prices

Regression results indicate that Opec is able to influence real oil prices via capacity utilization, by setting production quotas, and the degree to which Opec production exceeds these quotas. These variables reflect three decision variables; capacity, production, and quotas. Opec's ability to affect price via capacity and production is not surprising given its market share and control over marginal supply. Beyond capacity utilization, Opec production decisions affect prices via two variables, Quota and Cheat. The similarity in the regression coefficients associated with these values in Table 2 may undermine the importance of Opec quotas. If $\beta_2 = \beta_3$ in equation (1), the cheat and Quota variables simplify to production (the effect of changing quota is offset by its effect on Cheat). This would imply that Quota has no effect on real oil prices beyond Opec production. A similar interpretation applies to the corresponding elements of the cointegrating vector in equation (3).

To investigate the effect of Opec quotas on real oil prices, we test the restriction $\beta_2 = \beta_3$ in the DOLS estimate of equation (1) and the FIML estimate of equation (3). For equation (1), we fail to reject $\beta_2 = \beta_3$ [t(23) = 0.11 p > .91]. On the other hand, we strongly reject $\beta_2 = \beta_3$ in equation (3) [$\chi^2(1) = 22.03$, p < .001].

These confounding results may reflect uncertainty about the persistence of Opec decisions about quotas on price. In the short run, crude oil prices on the NYMEX and other exchanges respond to Opec decisions regarding quotas. This short-run effect provides an incentive for Opec as a group to cheat on its quota. If Opec quotas affect prices separately from production, Opec can raise prices by setting a low quota and overproducing relative to a scenario in which Opec sets its quota consistent with its intended level of production. But there probably is a time limit on this strategy. In the long run (the relationship estimated by the cointegrating relation), Opec announcements of small quotas relative to production should be recognized as a sham and should not boost prices beyond the level consistent with the fundamentals of the supply/demand balance.

Surprisingly, there is no evidence that real oil prices affect the degree to which Opec cheats on the quota. Restrictions on the VAR and the VECM do not allow us to reject the null hypothesis that real oil prices do not "Granger cause" the degree to which Opec cheats on its quota. This indicates that cheating by individual members of Opec may be prompted by reasons other than price. Possible incentives are discussed by Griffin and Xiong (1997).

Similarly, the analysis of causal order does not indicate that there is a feed back effect from real oil prices to Opec capacity utilization or Opec quotas. This implies that Opec considers objectives other than price when it sets production, capacity, and/or quotas. As described by Gately (1995), these objectives may include the net present value of oil revenues, a "fair share" of the market, and/or political goals.

Opec Supply Behavior

Estimates of β and γ from equation (4) and its variants are used to choose among competing models of Opec behavior. With regard to γ, Griffin (1985) estimates equation (4) with quarterly data from 1971:I through 1983:III and finds $\gamma < 0$ for five Opec nations. Jones (1990) updates these results with quarterly data that include a prolonged period of declining prices (1982:IV - 1988:IV). During this period, $\gamma < 0$ for only two Opec nations. Evidence for a negative relation between price and production is strengthened by Ramcharran (2002), who finds that $\gamma < 0$ for up to seven nations (depending on the threshold for statistical significance) in annual data from 1973-1997. Consistent with previous results, the estimate of γ in equation (6) is negative for eight of the ten Opec nations analyzed (Table 5).

The critical question is: what does the negative relation between price and production represent? Griffin (1985), Jones (1990), and Ramcharran (2002) offer no econometric evidence that prices affect Opec production. Nor is this causal ordering consistent with our analysis and that of Gulen (1997), which indicates that the negative relation may represent the opposite effect—that changes in Opec production affect price. That is, increases in Opec production reduce prices.

The negative effect of production on price generally is consistent with the analysis of causal order for individual Opec nations. For eight nations, there is little support for the notion that production responds to price (Table 5). That is, there is no evidence that Kuwait, Qatar, UAE, Libya, Algeria, Iran, Nigeria, or Indonesia reduce production in response to higher prices. There is evidence for a causal relation from price to production for Saudi Arabia and Venezuela at the five percent threshold and Algeria at the ten percent threshold. Algeria (and at times Venezuela) is considered a “price hawk” and a backward bending supply curve may be consistent with this preference. Alternatively, the target revenue models of Teece (1992) and Cremer and Salehi Isfahani (1989) imply that there should be a causal relation from price to production in nations constrained by their ability to absorb revenue. Constraints on absorptive capacity may be present in Algeria, but such constraints are unlikely in Saudi Arabia and Venezuela.

On the other hand, there is evidence that prices respond to production by individual Opec nations. For four nations, UAE, Algeria, Libya, and Venezuela, we reject the null hypothesis of no causal order from production to price at the five-percent threshold. If we use the ten-percent threshold, this list grows by another two nations, Saudi Arabia and Iran. The effect of production by Saudi Arabia on price is not surprising. Griffin and Xiong (1997) argue that Saudi Arabia maintains cartel discipline by producing more than its quota to punish cheaters. Similarly, Venezuela is a large producer and it ignored production quotas and challenged the Saudi leadership of Opec prior to the election of Hugo Chavez in 1999 (Fritseh, 1997; Colitt, 1998).

Together, these results indicate that interpreting econometric estimates for β and γ in equation (4) and its variants to test competing hypotheses about Opec behavior may be misleading. The negative value for γ may represent the negative relation between price and production in a cointegrating relation for price. We find little evidence that the negative relation between price and production for individual nations is part of a cointegrating relation for production. If equation (4) and its variants are not a cointegrating relation for production, γ and β cannot be used to test competing models for Opec behavior. This difficulty goes beyond the econometric complications that equations

(4) and its variants do not cointegrate and therefore, the estimates for β and γ are spurious.

The notion that any simple relation between price and production represents their relation in a cointegrating relation for price (and not production) is bolstered by the analysis of non-Opec nations. None of the six non-Opec nations show a statistically significant relation between price and production (Table 5). Nor is there any evidence that production and price cointegrate, either in the original specification used by Griffin (1985) or equation (6). We are not able to reject the null hypothesis that real oil prices do not “Granger cause” production for five nations at $p < .05$. These results are not surprising given the lack of a simple relation between real oil price and production in the US (Kaufmann and Cleveland, 2001). On the other hand, there is some evidence that production by Canada, Norway, and the UK “Granger cause” price. These three non-Opec nations increased production during the sample period. This allowed them to gain market share relative to Opec (given Opec’s roles as a residual supplier) and this suppressed prices by limiting gains in Opec capacity utilization. Again, the analysis of causal order between prices and production in non-Opec nations indicates that the negative relation between price and production should be interpreted as part of a cointegrating relation for price, not production.

OECD oil stocks

OECD stocks of crude oil have a negative effect on real oil prices. This effect creates the potential for an externality in private decisions about the optimal level for crude oil stocks. Most analyses of private stocking decisions focus on the role of stocks in reducing the impacts of a supply disruption (e.g. Bohi and Montgomery, 1982). But our results indicate that private stocks also have an important social benefit—they hold down prices. The regression coefficients in Table 2 indicate that a one-day increase in OECD stocks reduces real oil prices by \$0.88 - \$1.45 per barrel. The benefits of this reduction probably are not be captured fully by those who determine optimal stocks.

Over the last ten years, private firms have changed their willingness to hold stocks. Many firms have adopted just-in-time practices to reduce inventory costs. Estimates for the costs of storing oil vary but one engineering study indicates that the operating and management costs are equal to or less than \$0.25 per barrel per year (PD-KBB, 1998). These private costs are considerably smaller than the increase in real oil prices that are implied by the historical decline in OECD stocks. In the mid and late 1980's, days of forward consumption in OECD oil stocks fluctuated in the low 90's (Figure 2). This level dropped to the mid and low 80's by the mid and late 1990's. The reduction of nearly ten days of forward consumption implies an increase in real oil prices of several dollars per barrel.

These numbers seem to imply that recent inventory decisions have reduced the private costs of holding stocks but may have increased the social costs associated with higher oil prices. If correct, there may be a significant externality in private decisions about optimal stocks. That is, the optimal level of stocks may be greater than current holdings. To remedy this externality, OECD nations could consider subsidizing private firms for a portion of the costs of holding stocks. One justification for this subsidy would be that it could potentially contribute account for the social benefits of lower oil prices that are generated by higher stock levels. However, this policy implication does not follow from our analysis. First, increasing stocks would lower prices, which would increase demand. This increase in demand would raise prices and thereby offset some of the gains associated with higher stocks. Second, an incipient reduction in crude prices would have an adverse impact on production, thereby putting upward pressure on prices. Third, the cost/benefit ratio of efforts to internalize this externality probably are not symmetric. The cost of increasing storage capacity is two to three orders of magnitude greater than operating and management costs (PD-KBB, 1988). Low operating and management costs imply that effective policies should prevent the loss of existing capacity. But once capacity has been lost, the social benefits of lower oil prices associated with expanding storage capacity and stocks may be less than the private costs of this expansion.

V Concluding remarks

The results of this analysis indicate that Opec has considerable power over price via decisions about quotas, production, and operable capacity. Decisions about these variables by Opec and individual nations are relatively independent of real oil prices. Indeed, the negative relation between price and production should not be interpreted as evidence for a backward bending supply curve for Opec or most of its members. Without a cointegrating relation for production, analysts need to rethink the way in which they test competing hypotheses about Opec behavior. Unfortunately, this analysis says little about the nature of Opec (cartel, dominant producer and fringe, etc) or the factor(s) that Opec uses to set production, quotas, and/or capacity. To alleviate this shortcoming, ongoing research is directed towards identifying a cointegrating relation for oil production.

Another caveat with our results relates to how they can be used to predict future developments. In particular, the increasing importance of production from some non-Opec countries (most notably Russia) poses a potential threat to Opec's ability to influence the oil markets. While the latest evidence on the behavior of oil prices is consistent with the notion of maintained Opec market power, further research is needed to address this issue in more depth.

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Data Appendix

Data for oil prices, OECD oil stocks, and OECD demand for crude oil are obtained from the Monthly Energy Review (various years). Annual values for Opec capacity are obtained from Erik Kriel of the US DOE/EIA. Quarterly values for Opec capacity are interpolated by assuming a constant rate of change between annual observations. Data for Opec production are obtained from the OPEC Annual Statistical Bulletin (Figure 1).

Literature Cited

Akaike, H., 2nd International Symposium on Information Theory, P.N. Petrov and F. Csaki, Eds., Akadacemiai Kiadaco, 267-281, 1973.

Andrews, D.W.K. Tests for parameter instability and structural change with unknown change point. *Econometrica* 61(4): 821-856, 1993

Balabanoff, S. Oil futures prices and stock management: a cointegration analysis. *Energy Economics* 17(3): 205-210, 1995.

Belton, B. Opec can no longer make “people quake” *USA Today*, March, 27 1998.

Bohi, D.R. and W.D. Montgomery, *Oil Prices, Energy Security, and Import Policy,* Resources for the Future, Washington, DC., 1982.

Colitt, R, Venezuelan oil output above quotas, *Financial Times,* January 15, 1998.

Cremer, J. and D. Salehi-Isfahani, The rise and fall of oil prices: a competitive view. *Annals D'economie et De Statistique* 15/16: 427-454, 1989.

Dahl, C. and M. Yücel, Testing alternative hypotheses of oil production behavior. *The Energy Journal* 12(1): 117-138, 1991.

Dickey, D. A. and W.A. Fuller, Distribution of the estimators for auto regressive time series with a unit root. *Journal of American Statistics Association*. **7**: 427-431, 1979.

Engle, R.E. and Granger, C.W.J. Cointegration and error-correction: representation, estimation, and testing. *Econometrica* 55: 251-276, 1987.

Fritseh, P., Venezuela expands oil industry, irking others in Opec, *Wall Street Journal* August 14, 1997.

Gately, D. Strategies for Opec's pricing and output decisions. *The Energy Journal*, 16(3): 1-38, 1995

Gately, D. and J.F. Kyle. Strategies for OPEC's pricing decisions. *European Economic Review*. 10: 209-230, 1977.

Granger, C.W.J. Investigating causal relations by econometric models and cross spectral models. *Econometrica* 37: 424-438, 1969.

Granger, C. W. J. and P. Newbold, Spurious regressions in econometrics. *Journal of Econometrics, 2:111-120, 1974.*

Gregory, A.W. and B.E. Hansen, Residual-based tests for cointegration in models with regime shifts. *Journal of Econometrics* 70: 99-126, 1996.

Griffin, J.M. Opec Behavior: a test of alternative hypotheses. *American Economic Review,* 75(5): 954-963, 1985.

Griffin , J.M. and Xiong, The incentive to cheat: an empirical analysis of Opec. *Journal of Law and Economics* 40:289-316, 1997.

Gulen, S.G. Is Opec a cartel? Evidence from cointegration and causality tests. *The Energy Journal* 17(2): 43-57, 1997.

Hylleberg, S, R.F. Engle, L.W.J. Granger, and B.S. Yo, Seasonal integration and cointegration. *Journal of Econometrics* 44: 215-238, 1990.

Johansen, S., Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control*, 12: 231-254, 1988.

Johansen, S. and K. Juselius, Maximum likelihood estimation and inference on cointegration with application to the demand for money. *Oxford Bulletin of Economics and Statistics*, 52: 169-209, 1990.

Jones, C.T., Opec behavior under falling prices: implications for cartel stability, *The Energy Journal* 11(3): 117-129, 1990.

Kaufmann, R.K. A model of the world oil market for Project LINK: Integrating economics, geology, and politics. *Economic Modeling*. 12(2):165-178, 1995.

Kaufmann, R.K. and C.J. Cleveland. Oil production in the lower 48 states; Economic, geological, and institutional determinants. *The Energy Journal*. **22:**27-49, 2001

Mackinnon, J.A., Approximate asymptotic distribution functions for unit-root and cointegration tests. *Journal of Business and Economic Statistics* 12, 167-176, 1994.

Newey, W.K. and K.D. West. A simple positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica* 55, 703-708, 1987.

Ohanian, L.E. The spurious effects of unit roots on vector autoregressions. *Journal of Econometrics* 39:251-266, 1988.

Osterwald-Lennum, M., A note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics, *Oxford Bulletin of Economics and Statistics*, 54, 461-471, 1992.

PB-KBB Inc. Strategic oil storage concepts and costs for Asia Pacific region, Final draft report prepared for APERC (Asia Pacific Energy Research Center) October 30, 1998.

Powell, S.G. The target capacity-utilization model of Opec and the dynamics of the world oil market *The Energy Journal* 11(1): 27-63, 1990.

Ramcharran, H. Oil production responses to price changes: an empirical application of the competitive model to Opec and non-Opec countries, *Energy Economics* 24:97-106, 2002.

Sims, C., Macroeconomics and reality. *Econometrica* 48:1-49, 1980.

Stock, J.H., Asymptotic properties of least squares estimators of co-integrating vectors *Econometrica* 55,1035-1056, 1987.

Stock, J.H. and M.W. Watson, *Introduction to Econometrics*, Addison Wesley, Boston, 2003.

Stock, J.H. and M.W. Watson, A simple estimator of cointegrating vectors in higher order integrated systems *Econometrica* 61, 783-820, 1993.

Teece, D.J. Opec behavior: an alternative view in J.M. Griffin and D. Teece (eds). Opec Behavior and World Prices London: Allen and Unwin, 1992.

Toda, H.Y. and P.C.B. Phillips. The spurious effects of unit roots on exogeneity tests in vector autoregressions: an analytic study. *Journal of Econometrics* 59:229-255, 1993.

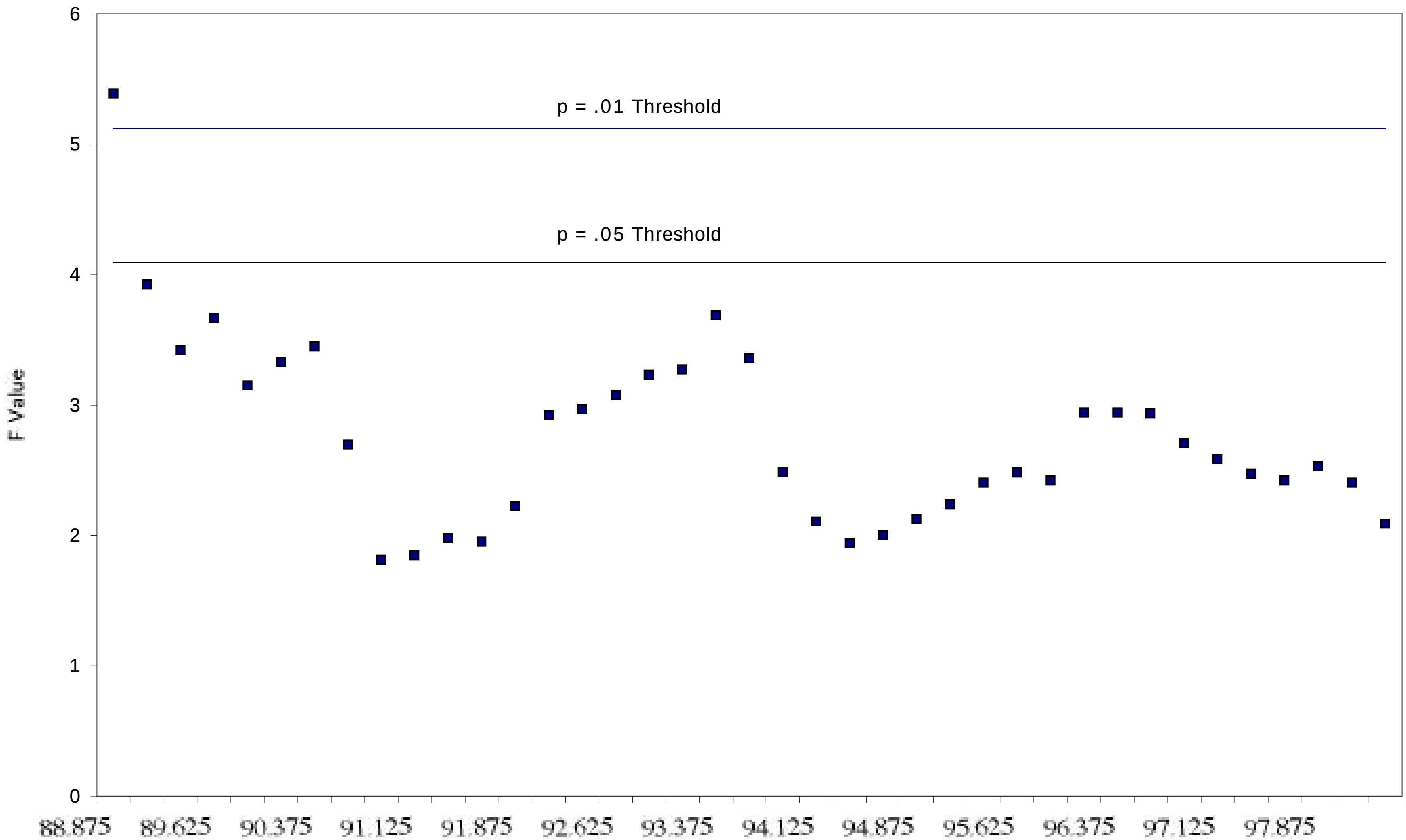
Figure Captions

Figure 1 The Quandt Likelihood ratio estimated from the OLS estimate of equation (1). The five percent (4.09) and 1 percent (5.12) critical values are obtained from an updated version of Andrews (1993) as reported in Stock and Watson (2002).

Figure 2 Opec decision variables. Opec capacity (dotted line), Opec production (dashed line), and Opec quotas (solid line).

Figure 3 OECD stocks of crude oil.

Figure 1



6
5
4
3
2
1
0
F Value
p = .01 Threshold
p = .05 Threshold
88.875
89.625
90.375
91.125
91.875
92.625
93.375
94.125
94.875
95.625
96.375
97.125
97.875

Figure 2

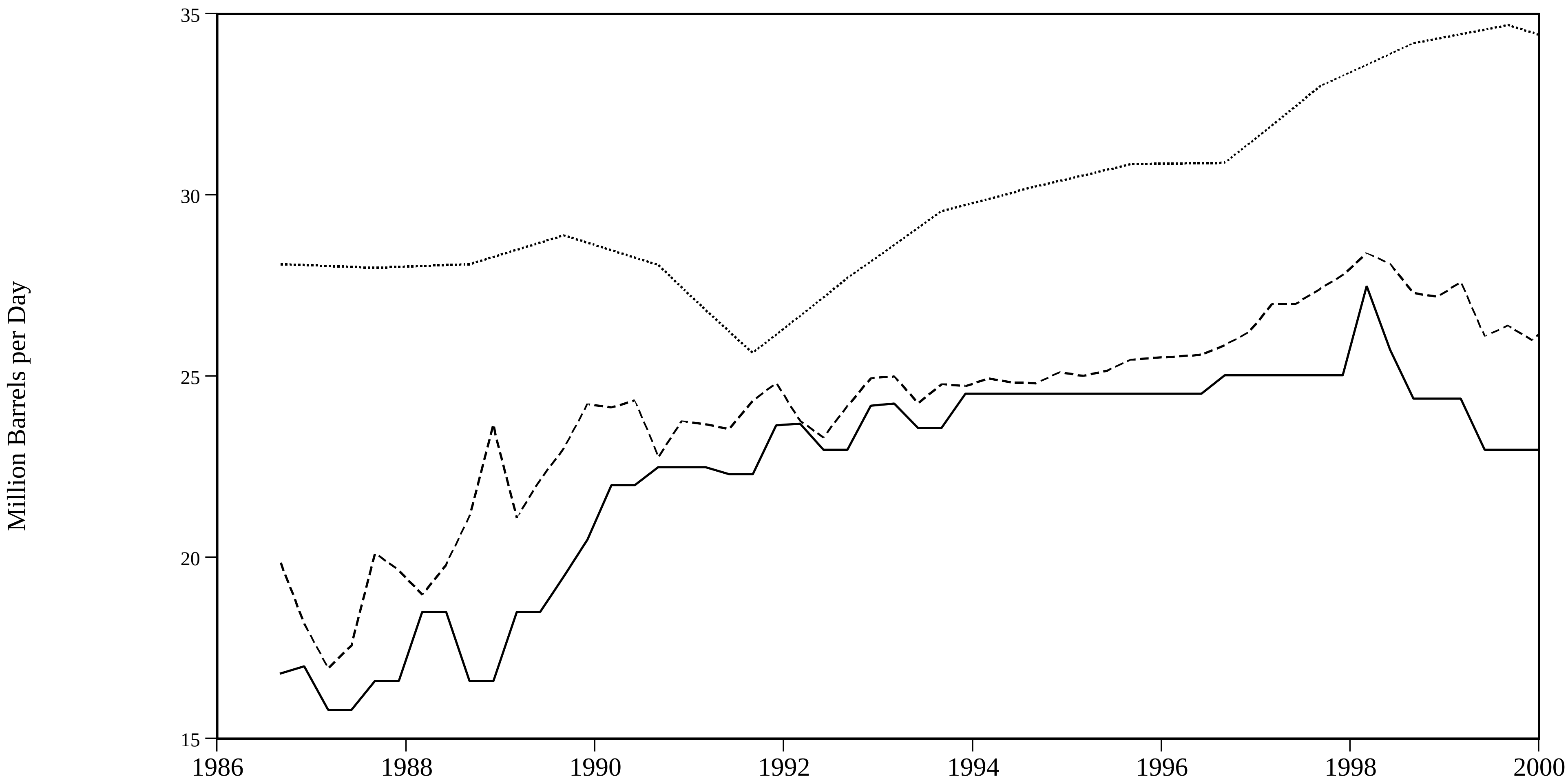

Million Barrels per Day
35
30
25
20
15
1986
1988
1990
1992
1994
1996
1998
2000

Figure 3

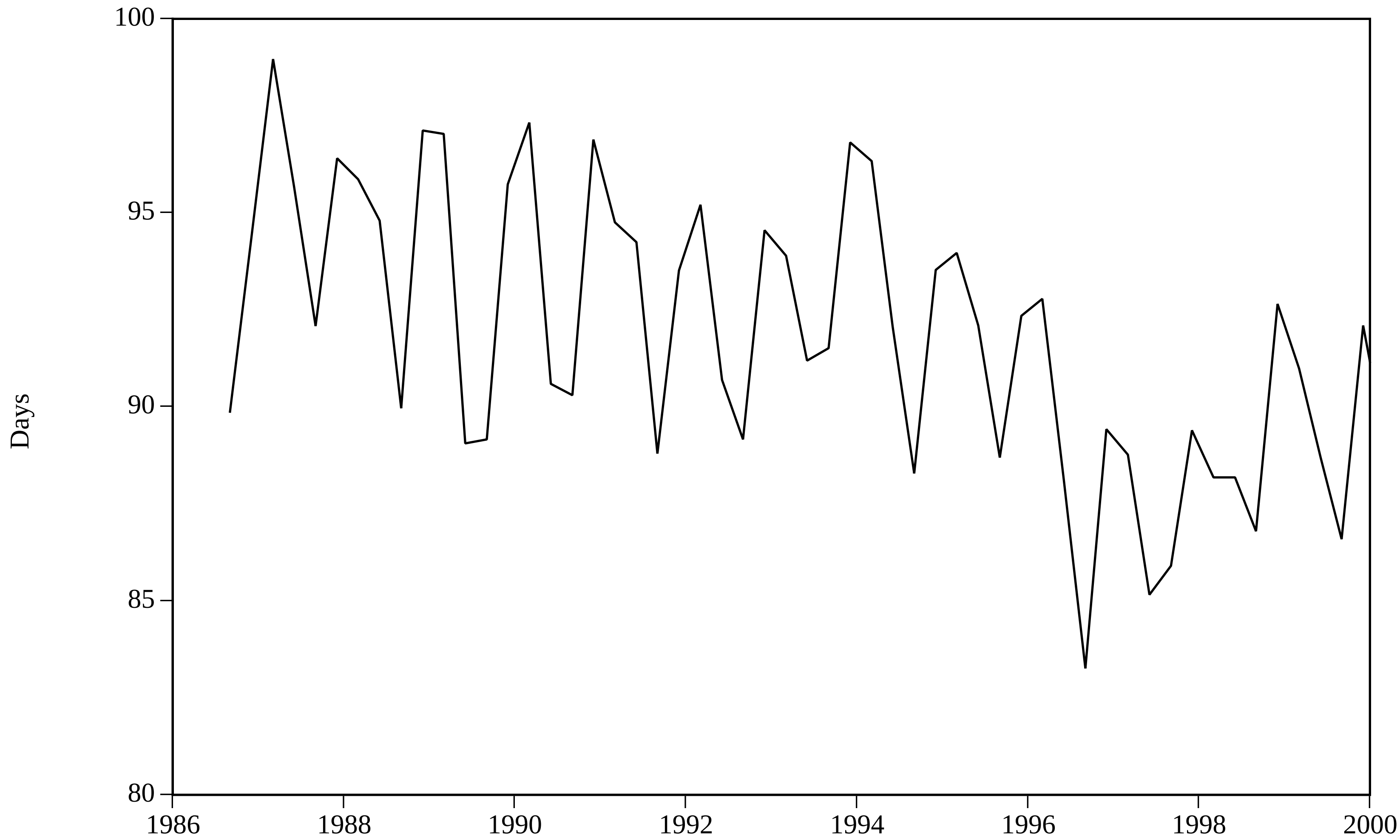


Table 1
HEGY statistics for annual and seasonal unit roots

	ADF	π_1	π_2	π_3	π_4	F $\pi_3 \cap \pi_4$
Univariate tests						
Price	-2.94	-3.83*	-7.98**	-5.21**	-2.02**	190.1**
Days	-2.45	-4.09*	-4.29**	-4.07*	-0.98	8.93**
Quota	-1.51	-2.31	-4.48**	-3.69*	-3.64**	16.87**
Cheat	-1.11	-2.16	-2.62	-3.19	-2.70*	8.34**
Caputil	-2.02	-3.21	-2.50	-2.32	-3.49**	9.02**
OLS Regression residuals						
Equation (1)	-4.51*	-4.13**	-3.63**	-4.90**	-0.61	12.22**
Equation (5)						
Saudi Arabia	-3.51	-2.40	-4.16**	-3.66*	-2.24*	10.86**
Kuwait	-4.19	-3.34	-5.24**	-1.95	-4.57**	14.47**
Qatar	-2.83	-3.19	-5.02**	-5.39**	-3.78**	34.29**
UAE	-3.01	-2.59	-5.87**	-5.95**	-2.60**	25.01**
Libya	-3.61	-1.80	-5.75**	-4.26*	0.10	9.12**
Algeria	-3.97	-2.62	-5.02**	-4.72**	-1.52+	13.82**
Iran	-3.54	-2.80	-5.73**	-5.73**	-1.73+	21.41**
Nigeria	-3.03	-2.82	-5.27**	-4.95**	-2.58**	19.55**
Indonesia	-4.32	-3.49	-4.52**	-5.77**	-1.24	17.42**
Venezuela	-2.74	-2.43	-5.55**	-4.62**	-1.59+	13.04**
Equation (7)						
Canada	-3.97	-0.78	-4.03	-3.55	02,36	10.61**
China	-3.93	-0.89	-3.20	-4.85**	-3.38	21.63**
Mexico	-2.96	-1.43	-3.65	-4.69**	-2.39	16.70**
Norway	-3.64	-2.18	-2.80	-4.15**	-3.44	16.80**
UK	-2.70	-1.99	-2.52	-1.54	-0.13	1.20
US	-4.14	-0.46	-1.83	-2.09	-3.33	8.16**

** Value exceeds $p < .01$, *$p < .05$, and + $p < .10$

Univariate HEGY tests inclue an intercept, time trend, and seasonal dummies HEGY test on OLS regression residuals do not include an intercept, time trend, or seasonal dummies. Significance levels from Hylleburg *et al*., (1990).

Univariate ADF test includes an interecept time trend, and seasonal dummies. ADF tests of cointegrating relation include a constant. Number of augmenting lags chosen using the Akaike information criterion (Akaike, 1973). Significance levels from Mackinnon (1994).

Table 2
Estimates for Price Equation

	OLS	DOLS	FIML
Cointegrating Relation			
Price			1.00 $\chi^2(1)$ 49.60
Constant	123.67 (6.52)	174.50 {3.63}	-105.96 $\chi^2(1)$ 17.96
Days	-0.992 (5.64)	-1.45 {3.35}	0.88 χ^2 (1) 15.70
Caputil	27.42 (3.64)	32.47 {2.30}	-38.25 χ^2 (1) 25.20
Cheat	-1.25 (3.39)	-2.00 {2.87}	0.572 χ^2 (1) 2.64
Quota	-1.62 (6.18)	-2.05 {2.85}	1.725 χ^2 (1) 25.49
Short run Dynamics			
Adjustment rate (α)	-0.504 (2.40)	-0.56 (2.11)	-0.79 (13.1)

t statistics in parenthesis
{} t statistics calculated using the Newey-West (1987) estimator
Chi square statistic tests exclusion from the cointegrating vector.
** Value exceeds $p < .01$, *$p < .05$, and + $p < .10$

Table 3
Tests of Causal Order on the VAR

	Independent Variable				
Dependent Variable	Price	Days	Caputil	Quota	Cheat
Price	F(4,28)= 14.77**	F(4,28)= 5.10**	F(4,28)= 5.89**	F(4,28)= 6.29**	F(4,28)= 2.13
Days	F(4,28)= 0.35	F(4,28)= 3.48*	F(4,28)= 0.59	F(4,28)= 1.27	F(4,28)= 0.60
Caputil	F(4,28)= 1.48	F(4,28)= 0.72	F(4,28)= 14.23**	F(4,28)= 0.70	F(4,28)= 1.67
Quota	F(4,28)= 1.40	F(4,28)= 0.61	F(4,28)= 0.74	F(4,28)= 7.14**	F(4,28)= 1.65
Cheat	F(4,28)= 0.97	F(4,28)= 0.66	F(4,28)= 0.40	F(4,28)= 0.52	F(4,28)= 1.20

Values represent F statistics
** Value exceeds p < .01, *p < .05, and + p < .10

Table 4
Exclusion tests on FIML

Variable	$\alpha = 0$
Price	$\chi^2(2)$ 53.22**
Days	$\chi^2(2)$ 0.03
Quota	$\chi^2(2)$ 4.51
Caputil	$\chi^2(2)$ 15.89**
Cheat	$\chi^2(2)$ 2.17

** Value exceeds p < .01, *p < .05, and + p < .10
ECM with four lags and two cointegrating relations

Table 5
The relation between price and production in Opec and Non-Opec nations

	γ	r^2	Prod → Price	Price → Prod
Opec (equation 6)				
Saudi Arabia	-0.074* (2.90)	0.98	F(4,24) = 2.49^{+}	F(4,24) = 2.93*
Kuwait	-0.40 (1.32)	0.80	F(4,24) = 0.16	F(4,24) = 1.05
Qatar	-1.00** (3.73)	0.98	F(4,24) = 1.01	F(4,24) = 0.30
United Arab Emirates	-0.136** (3.85)	0.94	F(4,24) = 1.73	F(4,24) = 1.44
Libya	-0.068* (2.96)	0.96	F(4,24) = 5.42**	F(4,24) = 2.00
Algeria	-0.107** (6.35)	0.91	F(3,30) = 9.34**	F(3,30) = 2.43^{+}
Iran	-0.165** (4.87)	0.94	F(4,24) = 2.66^{+}	F(4,24) = 1.40
Nigeria	-0.092** (3.06)	0.95	F(1,42) = 2.18	F(1,42) = 0.48
Indonesia	-0.035^{+} (1.94)	0.89	F(4,24) = 1.00	F(4,24) = 1.37
Venezuela	-0.065** (2.94)	0.98	F(4,24) = 5.53**	F(4,24) = 3.31*
Non-Opec Production (Equation 8)				
Canada	-0.031 (1.25)	0.92	F(4,24) = 4.30**	F(4,24) = 1.50
China	-0018 (1.16)	0.94	F(4,24) = 0.98	F(4,24) = 0.27
Mexico	-0.063^{+} (1.81)	0.74	F(4,24) = 0.93	F(1,42) = 0.97
Norway	0.118 (1.57)	0.96	F(4,24) = 2.76^{+}	F(4,24) = 2.84^{+}
UK	-0.129 (1.62)	0.74	F(4,24) = 3.69^{+}	F(4,24) = 1.01
US	-0.042^{+} (1.78)	0.94	F(4,24) = 1.03	F(4,24) = 3.79*

** Value exceeds p < .01, *p < .05, and + p < .10