A STUDY OF CARTEL STABILITY:

THE JOINT ECONOMIC COMMITTEE

1880 - 1886

by

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Abstract

This paper employs weekly time series data on the Joint Economic Committee railroad cartel from 1880 to 1886 to empirically test the proposition that observed price wars represented a switch from collusive to noncooperative behavior. An equilibrium model of dynamic oligopoly with assymetric firms, together with explicit functional form assumptions about costs and demand, determines the estimating equations and stochastic structure of the econometric model. This hypothesis is tested against the alternative that no switch took place, so that price and quantity movements were solely attributable to exogenous shifts in the demand and cost functions.

1. Introduction

Industrial organization economists have recognized for some time that the problem of distinguishing empirically between collusive and noncooperative behavior, in the absence of a "smoking gun," is a difficult one. A variety of approaches have been adopted in an attempt to surmount this problem. For instance, some researchers have studied industries characterized by homogeneity of product quality, with each firm facing the same production technology. They then assume that firms maintain constant conjectural variations, that is, that their rivals have reaction functions of constant slope. Within such a structure it is then possible to estimate the conjectural variation, and so infer the extent to which industry behavior is collusive. (See, for example, the work of Appelbaum [1], Gollop and Roberts [4], and Iwata [6].)

An alternative research methodology is to examine an industry both before and after it is affected by some structural change. Just and Chern [7] showed that the behavior of the tomato processing industry was consistent with oligopsonistic dominant firm price leadership, as they put it. They were able to identify noncompetitive behavior by studying the impact of a technological shock, namely the availability of a more productive harvesting machine. Somewhat analogously, Kosobud and Stokes [9] demonstrated that the market shares of oil producing countries were more stable after 1973, taking this as evidence that OPEC members were following a strategy of maintaining constant market shares, and hence were behaving collusively. (Osborne [11] has shown that in a world of complete information, such strategies will lead the participants of a cartel to the joint profit maximal outcome.)

Bresnahan [2], on the other hand, considered explicitly the implications of collusive as opposed to noncooperative behavior in an industry with product differentiation. Specific functional form assumptions allowed him to discriminate between the two. He then employed non-nested hypothesis tests to conclude that, in the American automobile industry, the 1954 and 1956 model years were collusive, but that there was a price war in 1955.

This paper adopts yet another research strategy, originally proposed in Green and Porter [5]. They considered an explicitly dynamic model in which the firms of an industry are faced with the problem of detecting and deterring cheating on an agreement. In particular, they assumed that firms observe only their own production level and the market price, but not the quantity produced by any other firm. (Their output is assumed to be of homogeneous quality, and so they face a common market price.) If the market demand curve has a stochastic component, an unexpectedly low price may signal either deviations from collusive output levels or a "downward" demand shock. Under these circumstances, participating firms can deter deviations by threatening to produce at Cournot levels for a period of fixed duration whenever they observe market price below some trigger price. A firm which considers a secret expansion of output above the collusive level must trade off immediate profit gains with the increased probability that the market price will fall below the trigger price, thereby increasing the likelihood of lower profits when the industry reverts to Cournot output levels. Their paper offers an explanation that what looks like collusive behavior at a point in time is actually the noncooperative outcome of a regularly repeated market game. For small enough discount rates, the joint profit maximizing output vector can be supported as a noncooperative equilibrium. Thus the results of Friedman [3] and Telser [14] extend to uncertain environments. In equilibrium, firms maximize expected discounted profits by producing at collusive output levels, so that any price wars which are observed should occur after unexpected drops in demand, rather than actual cheating by member firms. Thus price wars can be the occasional equilibrium outcome of a dynamic noncooperative market game.

Green and Porter go on to propose econometric techniques which employ aggregate time series price and quantity data for a particular industry, and which are designed to detect such an enforcement mechanism. The equilibrium conditions of the theoretical model determine the estimating equations and stochastic structure of the econometric model. This model exploits the fact that there will be periodic switches or reversions between the Cournot and collusive output levels when such a noncooperative equilibrium exists. These reversions serve to identify periods of collusive behavior in a simultaneous equation switching regressions model. The present paper applies this methodology to weekly data on the Joint Economic Committee, henceforward referred to as the JEC, from 1880 to 1886 inclusive. The JEC was a cartel which controlled eastbound rail transportation from Chicago to the eastern seaboard.

The JEC is described in more detail in the next section. In particular, the extent to which it corresponds with the paradigm presented above is documented. The econometric model is described in section 3, and its relationship to the theoretical model of Green and Porter outlined. The data set is summarized in section 4, as are the econometric results together with an interpretation of them in section 5. Section 6 provides some summary remarks.

2. A Description of the JEC

This section contains a description of the JEC, with emphasis on the period from 1880 to 1886. Readers who are interested in a more complete history should refer to MacAvoy [10] and Ulen [15]. Much of the material in this section is drawn from these studies.

The JEC was a cartel which controlled eastbound freight shipments from Chicago to the Atlantic seaboard in the 1880's. It was formed in June 1879 by an agreement of the railroads involved in the market. They intended it to be capable of preventing the problems which previous cartel structures had ineffectively dealt with. Until that time, many informal agreements had been repeatedly broken by secret cheating by member firms. The JEC, on the other hand, was explicitly designed so that internal enforcement mechanisms would deter defections, much as an external enforcement mechanism (e.g., a legal prohibition on cheating) might have done. The firms involved publicly acknowledged this agreement, as it preceded the passage of the Sherman Act (1890) and the formation of the Interstate Commerce Commission (1887). A separate agreement was reached for westbound shipments on the same railroad lines, primarily because of the essential physical differences of the products being transported.

Through shipments of grain accounted for 73 per cent of all dead freight tonnage handled by the JEC. The railroads also handled eastbound shipments of flour and provisions, but, with one minor expection in early 1882, the prices charged for transporting these commodities were tied to the grain rate. None of these commodities is easily perishable, so speed of delivery was probably not an important factor by which JEC members could have differentiated their products. Furthermore, while different railroads shipped grain to different port cities (for example, Baltimore and New York), most of the wheat handled by the cartel was subsequently exported overseas, and the rates charged by different firms adjusted to compensate for differences in ocean shipping rates. Thus, the assumption that a homogeneous good was sold seems to have been approximately satisifed. Also, attention can be focused on the movement of grain with little loss of generality.

Prices, rather than quantity, has typically been thought to be the strategic variable of firms in the rail-freight industry. In particular, the specification of Green and Porter that industry conduct during reversionary periods was Cournot might be considered unrealistic. Econometrically, it is not very difficult to modify the model so that firms revert from collusive to Bertrand behavior (as they would if they were price setters), rather than Cournot (in the case of quantity setting firms). This shall be demonstrated in the next section. If firms are price setters, then the inference problem they face in detecting cheating is quite similar to that originally posed by Stigler [13]. In the case of the JEC, the cartel agreement took the form of market share allotments rather than absolute amounts of quantities shipped. Firms set their rates individually, and the JEC office took weekly accounts so that each railroad could see the total amount transported. As we shall see, total demand was quite variable, and so the actual market share of any particular firm would depend on both the prices charged by all the firms as well as unpredictable stochastic forces. Thus the problem faced by the members of the JEC seems to be comparable to that posed by Green and Porter. For estimation purposes, the major change will be the specification of Bertrand behavior in reversionary periods.

In their model, Green and Porter explicitly rule out the possibility of entry into the market, since it is not clear how the cartel ought to react. In some industries, entry has triggered price wars, which may represent a predatory response on the part of the original firms, or just the equilibrium outcome of a more uncertain environment. In the case of the JEC, entry occurred twice between 1880 and 1886. It appears that the cartel passively accepted the entrants, allocating them market shares, and thereby allowing the collusive agreement to continue. The reason for this is probably that when a firm entered the rail freight industry in the late nineteenth century, it faced a "no-exit" constraint. To quote Ulen [15, pp. 70-71], "A railway, once built, no matter on what flimsy foundation, was determined in the eyes of the courts more than a mere commitment of private capital. The judicial view was that a transport firm was a social investment - no matter who paid for it - (necessitating) a public commitment to its continued operation." To put it briefly, bankrupt railroads were relieved of most of their fixed costs and instructed to cut prices to increase business. "There was no way, given the court enforced no-exit constraint, in which to prevent the successful entry of even the most ill-equipped competitor. In fact, by trying to bludgeon down a rival through lowering its rates, a healthy cartel risked driving the rival into bankruptcy, from which state it would be resurrected as an unbeatable competitor. For that reason the wisest policy was to admit a rival to the collusion or to buy him out before he completed construction [15, p. 74]." I have therefore chosen to deal with the actual entry which occurred during the sample period by appropriately modifying the nature of collusive and noncooperative outcomes, before and after entry, with the expectation that, ceteris paribus, reversionary periods should not have been precipitated by entry. Of course, entry to the industry may have increased the likelihood of future price wars.

The internal enforcement mechanism adopted by the JEC was a variant of a trigger price strategy. According to Ulen, there were several instances in which the cartel thought that cheating had occurred, cut prices for a time, and then returned to the collusive price. Price wars were not random, but precipitated by periods of slackened demand, which were presumably unpredictable, at least to some extent. On the other hand, the predictable fluctutations in demand that resulted from the annual opening and closing of the Great Lakes to shipping, which determined the degree of outside competition, did not disrupt industry conduct. Rather, rates adjusted systematically with the lake navigation season. Thus, in general, price wars were caused by unpredictable disturbances, as posited by Green and Porter, rather than by entry or predictable fluctuations in demand.

Lake steamers and sailships were the principal source of competition for the railroads, but at no point did they enter into an agreement with the JEC. In fact, there were several price wars in the lake shipping industry during the sample period, which were uncorrelated with the reversionary episodes of the JEC. Historically, railroads had captured a larger and larger share of the eastbound freight market, and this instability may have prevented a collusive agreement between the two modes of transport.

Therefore, the conduct of the JEC from 1880 to 1886 is largely consistent with the existence of a collusive equilibrium such as Green and Porter described. Whether the JEC participants explicitly used price as a monitoring variable, or whether they used some market share estimator or other monitoring variable prone to error because of demand shocks, the trigger price strategies we described are likely to account well for industry performance. In summary, the JEC appears to be an appropriate candidate for such an econometric study, with appropriate modifications.

3. The Econometric Model

This section is concerned with the possibility of estimating a model of the Nash equilibrium proposed by Green and Porter, suitably altered to reflect the structure of the JEC, using time series data on price and aggregate output levels for the JEC. A simultaneous equation switching regression model is proposed, in which the parameters of the demand and cost functions are estimated, and in which the regime probabilities are unknown, but endogenously predicted.

I wish to estimate a simultaneous system of two equations corresponding to a demand equation, which will be invariant across regimes, and a pseudo-supply function, which will vary across regimes to reflect collusive or Bertrand behavior on the part of firms. These two equations are now specified.

Denote the market price in period t by p_t. Then the total quantity demanded is assumed to be a loglinear function of price,

$$\text{(1)} \qquad \log Q_t = \alpha_0 + \alpha_1 \log p_t + \alpha_2 L_t + U_{1t}$$

where L_t is a dummy variable equal to one if the Great Lakes were open to navigation, and $\{U_{11}, U_{12}, \ldots U_{1T}\}$ is a sequence of independently distributed normal variables with zero mean and variance σ_1^2. Here α_1 is the price elasticity of demand, and presumably negative. Also α_2 should be negative, reflecting a decrease in demand when the lake steamers were operating. (I also experimented with another additive term, $\alpha_3 L_t \log p_t$, where α_3 should be negative if demand was more elastic when the lakes were open.)

The N active firms in the industry are assumed to be asymmetric, in that they each face a different cost function. The cost of producing output q_{it} for firm i in period t is given by

$$C_i(q_{it}) = a_i q_{it}^{\delta} + F_i, \qquad \text{for } i = 1,\ldots N,$$

where δ, the (constant) elasticity of variable costs with respect to output, must exceed one if an equilibrium is to exist. Here a_i is a firm-specific shift parameter, and F_i the fixed cost faced by firm i. These fixed costs are assumed to be small enough that firms have positive discounted expected profits in equilibrium. Marginal costs faced by firm i at output q_{it} are then

$$MC_i(q_{it}) = a_i \delta q_{it}^{\delta - 1}, \qquad \text{for } i = 1,\ldots N.$$

Under this specification, different firms can face average cost functions of radically different shape.

Since the products provided by these firms are of approximately homogeneous quality, all firms will charge equal prices in equilibrium. If firms are behaving noncooperatively in each period, they will price at marginal costs as Bertrand predicted, so that in reversionary periods firm i will produce q_{it} so that

$$p_t = MC_i(q_{it}), \qquad \text{for } i = 1,\ldots N.$$

If instead firms behave collusively, so that they maximize single period expected joint profits in the industry, then market price and the output level of each firm i must satisfy

$$p_t(1 + 1/\alpha_1) = MC_i(q_{it}).$$

If firms produce at Cournot output levels, then

$$p_t(1 + s_{it}/\alpha_1) = MC_i(q_{it})$$

where $s_{it} = q_{it}/Q_t$, the market share of firm i in period t.

For estimation purposes, I employ aggregate data. Thus these three behavioral relationships must be aggregated to give a market supply relationship in each case. In order to do this, each of the individual supply equations are weighted by market shares in time t, s_{it}, and added up. Then we get the supply relationship

$$p_t = \sum_i s_{it} MC_i(q_{it})$$

when behavior is Bertrand, i.e., Nash in prices,

$$p_t(1 + 1/\alpha_1) = \sum_i s_{it} MC_i(q_{it})$$

when firms behave collusively, or as a monopolist would, and

$$p_t(1 + H_t/\alpha_1) = \sum_i s_{it} MC_i(q_{it})$$

if behavior is Cournot, i.e., Nash in quantities. Here $H_t = \sum_i s_{it}{}^2$, the value of the Herfindahl index in period t.

In an appendix to their paper, Green and Porter [5] show that, given these functional forms for the market demand and cost functions, the market share of firm i in period t will be

$$s_{it} = \frac{a_i{}^{1/(1-\delta)}}{\sum_j a_j{}^{1/(1-\delta)}}$$

$$\equiv s_i.$$

in each of the three cases above. Thus the market share of each firm will be constant over time and invariant across changes in industry conduct.

Note that the market share of firm i is a decreasing function of a_i. The higher the value of the firm-specific marginal cost shift parameter, the lower is the market share of firm i.

The right hand side of each of the supply relationships can now be written as

$$\begin{aligned}\textstyle\sum_i s_{it} MC_i(q_{it}) &= \textstyle\sum_i s_i a_i \delta q_{it}^{\delta - 1} \\ &= \delta Q_t^{\delta - 1} \textstyle\sum_i a_i s_i^{\delta} \\ &= DQ_t^{\delta - 1},\end{aligned}$$

where the second equality follows from $q_{it} = s_i Q_t$, and where

$$D = \delta \left(\textstyle\sum_i a_i^{1/1-\delta} \right)^{1-\delta}.$$

Note that D depends only on the parameters of the cost functions of the firms. Thus the supply relationship is

$$p_t = DQ_t^{\delta - 1}$$

if firms have Bertrand behavior,

$$p_t(1 + 1/\alpha_1) = DQ_t^{\delta - 1}$$

if firms behave perfectly collusively, and

$$p_t(1 + H/\alpha_1) = DQ_t^{\delta - 1}$$

when they have Cournot conjectures. Here $H = \sum_i s_i^2$, so that the Herfindahl index is invariant across time if the number of firms remains unchanged.

Suppose I_t is an indicator variable which equals one when the industry is in a cooperative regime, and equals zero when the industry witnesses a reversionary episode. Then the supply relationship of the industry is given by

(2) $\log p_t = \beta_o + \beta_1 \log Q_t + \beta_2 S_t + \beta_3 I_t + U_{2t}$.

Since reversionary periods are Bertrand, $\beta_o = \log D$ and $\beta_1 = \delta - 1$. Since δ is assumed to be greater than one, β_1 should be positive. Here S_t is a vector of structural dummies which reflect entry and acquisitions in the industry. Recall that, for the JEC, entry does not seem to have caused reversions to noncooperative behavior. Then entry should not result in a regime change, only a shift in the parameter D. Also, $\{U_{21},...U_{2T}\}$ is assumed to be a sequence of independent normal variables, with mean zero, variance σ_2^2 and $\mathrm{COV}[U_{1t}, U_{2t}] = \sigma_{12}$.

If firms behaved in cooperative periods so as to maximize single period expected joint net returns, then β_3 would equal $\log(\alpha_1/(1 + \alpha_1))$. However, as I have shown elsewhere (Porter [12]), there is some reason to believe that if a cartel selects an optimal trigger price strategy, output in cooperative periods will exceed perfectly collusive levels. In the context of an industry with symmetric quantity-setting firms, linear demand and cost curves and no specified distribution of the demand shocks, I showed that the output in cooperative periods will exceed perfectly collusive levels when there is uncertainty, so that firms cannot infer from market price whether someone has cheated. In such a noncooperative equilibrium, the marginal gains from cheating in cooperative periods are exactly offset by the marginal losses implicit in the increased probability of an industry reversion to Cournot behavior. The marginal gains from cheating increase as output in cooperative periods decreases towards perfectly collusive levels, and so expected marginal losses must be increased by increasing the trigger price or the length of reversionary episodes. Expected discounted industry profits will be maximized at output levels between the Cournot and perfectly collusive quantities, given the constraint that firms have no incentive to cheat (i.e., expand output) in cooperative periods. While the industry structure described in this paper differs from that of Porter [12], there is some reason to suspect that the same sort of equilibrium will result. Thus the value of β_3 will not be restricted to equal $\log(\alpha_1/(1+\alpha_1))$, but instead will be estimated independently. Since market price should be higher in cooperative periods, β_3 should be positive but less than $\log(\alpha_1/1+\alpha_1)$. We can test whether cooperative periods involve significantly higher prices than reversionary periods by employing a one-tailed test of whether β_3 equals zero, versus $\beta_3 > 0$.

If the sequence $\{I_1,\ldots I_T\}$ is known, then the estimation of the parameters of the demand and supply functions is straightforward, as two stage least squares can be employed to obtain consistent estimates. If instead I_t is unknown, but assumed to be governed by the binomial distribution

$$(3) \qquad I_t = \begin{cases} 1 \text{ with probability } \lambda \\ 0 \text{ with probability } 1-\lambda. \end{cases}$$

then we have a simultaneous equations switching regression problem, where the "switch" is reflected solely by the constant term in the supply function. As is demonstrated in Green and Porter [5], the parameters of the demand and supply functions, as well as the switch probability λ, can be estimated by appropriately generalizing a technique first proposed by Kiefer [8], which adapts the E-M algorithm to models of this sort.

We can summarize equations (1) and (2) by writing

(4) $$By_t = \Gamma X_t + \Delta I_t + U_t$$

where

$$y_t = \begin{pmatrix} \log Q_t \\ \log p_t \end{pmatrix}, \quad X_t = \begin{pmatrix} 1 \\ L_t \\ S_t \end{pmatrix}, \quad U_t = \begin{pmatrix} U_{1t} \\ U_{2t} \end{pmatrix}$$

and where

$$B = \begin{pmatrix} 1 & -\alpha_1 \\ -\beta_1 & 1 \end{pmatrix}, \quad \Delta = \begin{pmatrix} 0 \\ \beta_3 \end{pmatrix}, \text{ and}$$

$$\Gamma = \begin{pmatrix} \alpha_o & \alpha_2 & 0 \\ \beta_o & 0 & \beta_2 \end{pmatrix}.$$

Here U_t is identically and independently distributed $N(0, \Sigma)$ where

$$\Sigma = \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{pmatrix}.$$

The probability density function of y_t, given I_t, is then

$$h(y_t|I_t) = (2\pi)^{-1}|\Sigma|^{-\frac{1}{2}}||B||\exp\{-\tfrac{1}{2}(By_t - \Gamma X_t - \Delta I_t)'\Sigma^{-1}(By_t - \Gamma X_t - \Delta I_t)\}$$

and the likelihood function, if there are T observations, is

$$L(I_1,\ldots,I_T) = \prod_{t=1}^{T} h(y_t|I_t).$$

If the $\{I_t\}$ sequence is known, then we can obtain estimates of B, Γ, Δ and Σ by maximizing $L(I_1,\ldots I_T)$. When the $\{I_t\}$ series is unknown and governed by equation (3), then the probability density function of y_t is given by

$$f(y_t) = (2\pi)^{-1}|\Sigma|^{-\frac{1}{2}}||B|| \times$$
$$[\lambda \exp\{-\tfrac{1}{2}(By_t - \Gamma X_t - \Delta)'\Sigma^{-1}(By_t - \Gamma X_t - \Delta)\}$$
$$+(1-\lambda)\exp\{-\tfrac{1}{2}(By_t - \Gamma X_t)'\Sigma^{-1}(By_t - \Gamma X_t)\}]$$

and the likelihood function by

$$(5) \qquad \pounds = \prod_{t=1}^{T} f(y_t).$$

Given an initial estimate of the regime classification sequence, say $\{w_1^o,\ldots,w_T^o\}$, where w_t^o is an estimate of $\Pr\{I_t = 1\}$, we can obtain an initial estimate of λ by using

$$\lambda^\circ = \sum_t w_t^\circ/T,$$

and initial estimates of Δ, Σ, B and Γ by maximizing $L(w_1^\circ,\ldots,w_T^\circ)$. Denote these estimates by $\Omega^\circ = (\Delta^\circ, \Sigma^\circ, B^\circ, \Gamma^\circ)$. Kiefer's algorithm then updates the w_t° series by Bayes' rule, so that

$$w_t^1 = \Pr\{I_t = 1|y_t, X_t, \Omega^\circ, \lambda^\circ\}$$
$$= \frac{\lambda^\circ h(y_t|X_t, \Omega^\circ, I_t = 1)}{\lambda^\circ h(y_t|X_t, \Omega^\circ, I_t = 1) + (1-\lambda^\circ)h(y_t|X_t, \Omega^\circ, I_t = 0)}$$

With the assumption that the covariance matrix Σ is constant across regimes, this expression is linear logistic in y_t, as is demonstrated in an appendix to Green and Porter [5]. Thus w_t^1 will be in the unit interval.

Given the new regime classification series $\{w_1^1,\ldots w_T^1\}$, new estimates of $(\Delta, \Sigma, B, \Gamma)$, say Ω^1, can be obtained by maximizing $L(w_1^1,\ldots w_T^1)$ with respect to Ω. Our new estimate of λ will be $\lambda^1 = \sum_t w_t^1/T$. This iterative procedure is continued until covergence occurs, say at $(\hat{w}_1,\ldots, \hat{w}_T)$, $\hat{\lambda} = \sum_t \hat{w}_t/T$, and $\hat{\Omega}$. The stopping criterion I employed was that the correlation between the estimated w_t sequences of two successive iterations exceed 0.999. As Kiefer shows, $\hat{\lambda}$ and $\hat{\Omega}$ will be the maximum likelihood estimates of λ and Ω. Thus $\hat{\lambda}$ and $\hat{\Omega}$ maxmize the likelihood function $\pounds$ of equation (5). (This is generally true for the E-M algorithm.)

The Kiefer estimation scheme does not constrain the estimated $\hat{w}_t$ series to follow any particular process. Instead, the $\hat{w}_t$ probabilities are estimated independently by the Bayesian updating formula. If trigger price strategies of the sort described by Green and Porter actually occur, then the $\hat{w}_t$ sequence should follow a Markov process of order equal to the length of reversionary periods. Rather than attempt to estimate the $\hat{w}_t$ sequence subject to a constraint of this sort, which would be relatively difficult, I have chosen to employ Kiefer's technique, and then see if the predicted $\hat{w}_t$ sequence is consistent with the existence of trigger price strategies. If so, these probabilities should fall sharply from high to low values in some periods, and vice versa after a "suitable" length of time, where this length of time should be roughly constant across episodes.

Once estimation is completed, several other types of tests can be carried out. From $(\hat{w}_1 \ldots \hat{w}_T)$ we will have estimated the probability of any particular observation belonging to either regime. Their average, $\hat{\lambda}$, is an estimate of the proportion of the total sample when firms behaved cooperatively. At a more basic level, likelihood ratio tests can be used to determine whether or not structural change has in fact occured. The natural alternative hypothesis is that only cooperative or noncooperative behavior is observed, but not both. The value of the likelihood function given the Kiefer estimation technique can be compared to that when

$\pounds$ is maximized subject to the constraint that $\Delta = 0$, (or, equivalently, when $L(I_1,\ldots,I_T)$ is maximized for $I_t = 0$ or 1 for all t).

In order to see how sensitive the estimation scheme is to the specified functional forms, I also estimated the model with a linear specification of equation (1) and (2), that is where $y_t' = [Q_t \; p_t]$.

4. The Data

A principle function of the JEC was information gathering and dissemination to member firms. Weekly accounts were kept in an effort to keep members abreast of developments in the industry. In this section I document the data set which is employed in this study, and mention some of its features. A list of variables is contained in Table 1, and TQG and GR series are depicted in Figures 1 and 2, respectively. Summary statistics for GR, TQG, PO and LAKES are provided in Table 2.

The quantity variable, TQG, is a reasonably accurate measure of the total tonnage of grain shipped by the members of the cartel. It changed dramatically over the sample period, but does not appear to follow any significant trend.

The price variable, GR, is somewhat suspect, however. The JEC provided it as an index of prices charged by member firms. There is some reason to expect that secret price cuts would not be reflected by this index, and so any price wars preciptated by secret price cutting may have been recorded with a lag, if at all. On the other hand, the existence of this sort of information structure is necessary if an enforcement mechanism involving reversions to noncooperative behavior, or price wars, is witnessed. It is of crucial importance that firms monitor some variable (in this case their own market share) which imperfectly reflects the actions of other firms. Here firms know what prices they charge their own customers, but the GR series would not be of much use in determining whether other firms were secretly cutting price.

The LAKES dummy variable is equal to one when the Great Lakes are open to navigation. While this variable accurately documents when the JEC faced its main source of competition, it would be preferable if the prices charged by the lake steamers has also been used in the econometric work. Unfortunately, this series was not available. As the mean of the LAKES series shows, the lakes were open in 57% of the sample weeks.

The PO series equals one unless the Railway Review, a trade magazine, reported that a price war was occurring. This series concurred with the reports of the Chicago Tribune and other accounts in this period. The PN series is the estimated w_t sequence, measuring the probability that industry conduct in period t is cooperative, which arises from Kiefer's estimation method, and which should mirror the PO series if the latter is at all accurate. One reason for estimating a PN series is that PO, gathered by Ulen [15], conflicts sharply with an index of cartel adherence created by MacAvoy [10].

The various DM dummy variables proxy structural change caused by entry, acquisitions or additions to existing networks. In each case, structural change is presumed to result in a once-and-for all shift in the constant term of the supply relationship, consistent with the algebra of the previous section.

Finally, I also employed dummy variables to capture seasonal aspects of market demand and supply. Each year was segmented into thirteen four-week segments, and so twelve "monthly" dummies entered both the demand the supply equations.

Table 3 gives the sample correlation of the GR, TQG, LAKES and PO series. The LAKES series is negatively correlated with both GR and TQG, which is consistent with downward shifts of the industry demand curve when the lakes were open. Consistent with an upward shift of the pseudo-supply curve in cooperative periods, PO is positively correlated with GR and negatively correlated with TQG. Finally, PO and LAKES are virtually independent, which accords with the previous assertion that systematic demand shifts should not cause price wars.

5. Results and Interpretation

This section contains an interpretive discussion of the econometric results, together with the outcomes of the tests described in section 3.

The regression coefficients obtained when two-stage least squares are applied to the system of equations (4), taking the PO series to be an accurate classification of regimes, are displayed in Table 4. Estimated t-statistics are enclosed in parentheses, and both single equation R^2 statistics and standard errors of the regression (s) are shown. Generally speaking, all variables have coefficients of the anticipated sign, significantly different from zero, but the "fits" are not particularly good. This is true of both the linear and loglinear regression models.

Since the two estimated systems are quite similar, I shall confine my attention to the loglinear equations. (The estimated elasticities are identical for practical purposes.) In the demand equation, the predicted quantity demanded is much lower when the lakes were open. The demand elasticity is negative, as expected, and less than one in absolute value, indicating that a price increase would be expected to increase total industry revenues. Since I do not have access to any cost data, it is impossible to say what the profit-maximizing price is, but it is possible to identify revenue-increasing changes in price. Recall that all the estimated demand equations will be marred by a missing data problem;

namely, the absence of a price series for lake transportation.

The supply equation is also fairly sensible. Price was significantly higher in cooperative periods. The predicted price of suppliers is an increasing function of quantity shipped, but the elasticity is of minor magnitude and only significantly different from zero at a 15 percent significance level. Given the presumed cost and demand functions, this might be taken as evidence of weak diseconomies of scale, at least locally. The coefficients of the structural dummies are also reasonable. Entry led to a fall in market price, *ceteris paribus*, as evidenced by the fact that the coefficient of DM1 is negative, and that of DM3 is less than that of DM2. Similarly, the acquisition of the Chicago and Atlantic line by the New York Central seems to have led to an increase in predicted price. I have no *a priori* reason to predict what the impact on market price of an addition to an existing network might be, and so no way of determining whether the relative magnitude of the coefficient of DM2 is sensible.

The fact that the coefficient of GR in the loglinear demand equation is less than one in absolute value implies that the marginal revenue associated with the industry demand curve is negative. This fact is not consistent with single period profit maximization, which stipulates that industry marginal revenue equal a weighted average of the marginal costs of individual firms, a positive number. To put this another way, if cooperative periods involved jointly maximizing behavior, then the coefficient of PO, β_3, should satisfy

$$\beta_3 = -\log(1 + 1/\alpha_1)$$

where α_1 is the price elasticity of demand. But this equation cannot be true if α_1 is negative and smaller than one in absolute value. If instead one supposed that

$$\beta_3 = - \log(1 + \Theta/\alpha_1)$$

for some constant Θ, then the estimated values of β_3 and α_1 imply that Θ equals 0.246. Since the implied magnitude of Θ is comparable to the Herfindahl index for the JEC in the sample period, then output levels in cooperative periods are similar to those implied by static Cournot behavior, given the that behavior is Bertrand during noncooperative episodes.

Table 5 displays the results of applying Kiefer's iterative technique. The coefficient attributed to PN is the estimate of β_3, i.e., the difference between the intercept of the supply relationship in cooperative and noncooperative periods. The obvious difference between the results of Tables 4 and 5 is that measures of goodness of fit of the supply equation are dramatically better in Table 5. Again, since the results of the linear system are analogous to those of the loglinear system, I shall focus my attention on the latter.

For practical purposes, the demand equations of Tables 4 and 5 are identical, as the estimated coefficients are not significantly different. The real differences are reflected in the supply relationships. The coefficients of the structural dummies continue to be reasonable, with the exception of that of DM4, in that it is less than the coefficient of DM3. (This is not true of the linear system.) However, the estimated standard error of the coefficient of DM4 is much larger (about 2.5 times the size of that of DM3), so that more plausible relative parameter values are not ruled out. The coefficient attributed to the PN series, β_3, is larger in Table 5, and with about half the standard error. If we again assume that

$$\beta_3 = - \log(1 + \Theta/\alpha_1)$$

for some constant Θ, then the value of Θ implied by the estimtes of β_3 and α_1 is 0.336. This is also roughly consistent with Cournot behavior in cooperative periods. The witnessing of approximately Cournot behavior is by itself of no special significance; what matters is that cooperative period prices exceed those implied by competitive price-setting, but are less than those consistent with static joint profit maximizing. Finally, the estimated TQG coefficient is closer to zero, but still greater than zero at about a 15 percent significance level.

If we set all explanatory variables equal to their sample mean, with the exception of the LAKES and PN dummy variables, then the structural system parameter estimates of the loglinear model, displayed in Table 5, imply the following reduced form estimates:

$$\log GR = -1.788 + (0.5077)PN - (0.0366)\ LAKES$$

$$\log TQG = 10.56 - (0.4060)PN - (0.4009)\ LAKES$$

Thus, in equilibrium, price is higher and quantity shipped lower in cooperative periods, and both endogenous variables lower when the lakes are open to navigation. To get an idea of the relative magnitudes involved, refer to Table 6. Price was 66 percent higher in cooperative periods (about 11 cents per 100 pounds), and quantity 33 percent lower. Similarly, price was 4.5 percent lower when the lakes were open (about half a cent per 100 pounds), and quantity 33 percent lower. The total revenue figure is five times the product of GR and TQG, and so in dollars (5 × ¢ per 100 lbs. × tons). Thus the cartel as a whole could expect to earn 11 percent higher revenues in cooperative periods, a difference of about $3000 per week. (Recall that these are 1880 dollars.) This is the revenue earned on grain shipments, which represented between 70 and 80 percent of total revenues from eastbound freight shipments by the JEC. Finally, note that revenues were about 35 percent lower when the lakes were navigable.

The PO and PN series can be compared by referring to Table 7, which shows when noncooperative episodes were predicted by the two series. If the PN series was approximately equal to zero for seven weeks in a row, then equal to .4 for one week and about equal to one the week after that, then I calculated the noncooperative period to be 7.4 weeks long. Both series are similar to the extent that noncooperative periods averaged about 10 weeks in duration, and seem to have occurred mostly in 1881, 1884 and 1885. In several instances, PO reflects a price war before PN, and both switch back to unity together, which is consistent with GR not picking up secret price cuts. For either series, a regression of price war length on the realization of the demand equation residual error term in the period prior to the beginning of the episode had little predictive power. Of course, the demand equation is marred by a missing variable problem (namely, the price charged by lake steamers), so there is not much reason to think that the demand residuals would accurately reflect unexpected disturbances. (Some people have suggested that optimal price war length might depend on the magnitude of the demand stock.) The 1881 and 1884 incidents both began about 40 weeks after the entry of the Grand Trunk and the Chicago and Atlantic, respectively. While entry may not have immediately caused reversion to noncooperative behavior, it is quite plausible that it increased the probability of its incidence in the future, as cartel enforcement problems typically increase with the number of participating firms. The final summary section contains further discussion of what may have caused the observed price wars, and to what extent their incidence is consistent with the Green and Porter paradigm.

As I mentioned before, the PO series collected by Ulen [15], differs markedly from an index of cartel non-adherence created by MacAvoy [10]. These series, as well as PN, are summarized in Table 8. MacAvoy had monthly data on market shares for each firm in the JEC. To get his non-adherence series, for each firm he regressed actual market share on a constant and a time trend. He then added up, for each year, the number of months in which the absolute value of the residual errors in these regressions exceeded the standard error of their respective regression equation. MacAvoy claimed that the greater the total number of months in which these residuals were too large, the more unstable the cartel, and hence the greater the incidence of noncooperative behavior. Thus his results purport to show that 1880 was the year of the greatest amount of non-adherence to the cartel agreement. (Figure 1 shows that this is when GR was at its highest levels.) The "Ulen" and "Estimated" columns show the fraction of weeks in each year in which PO and PN were equal to zero, respectively.

Since the PN series was in no way constrained to resemble PO, it is evident that PN supports the documentation of the <u>Railway Review</u> and <u>Chicago Tribune</u>, rather than MacAvoy's results. MacAvoy's procedure is probably faulty on two counts. First, he could have used a lot more information than deviations of market shares from time trends to detect incidents of noncooperative behavior. Secondly, it is not at all evident that instability of market shares necessarily implies the existence of noncooperative behavior. The demand and cost functions of section 3 are an example of a system in which market shares are invariant across different behavioral regimes.

To conclude this section, I consider the statistical evidence that switches actually occurred and were significant. First, the coefficients of PO in Table 4 and those attributed to PN in Table 5 are significantly greater than zero, so that periods of noncooperation involved a significantly higher price. The magnitude of the difference is indicated in Table 6.

Secondly, likelihood ratio tests can be used to discriminate between nested models. The estimates of Table 4 cannot be compared to those of Table 5 by this method, although the much better fit in the supply equations of Table 5 would incline me to favor the estimates obtained when Kiefer's algorithm is employed. Likelihood ratios can be used to compare the linear and loglinear models for the two estimation schemes, and to discriminate between models in which switching occurs and those in which it does not.

Suppose that L_1 is the maximized value of the loglikelihood function under the loglinear specification of Table 4, and L_0 the value under the linear specification of that table. If $(\hat{B}_1, \hat{\Sigma}_1)$ and $(\hat{B}_0, \hat{\Sigma}_0)$ are the respective two-stage least squares estimates of B and Σ, then the difference between L_1 and L_0 is given by

$$L_1 - L_0 = \log\{||\hat{B}_1|| - \tfrac{1}{2}|\hat{\Sigma}_1|) - (||\hat{B}_0|| - \tfrac{1}{2}|\hat{\Sigma}_0|) - T\tilde{Q}G - \tilde{G}R\}$$

where $\tilde{X}$ is the geometric mean of the sample values of the variable X. From our estimates, $(L_1 - L_0)$ is equal to 0.0341. Thus the loglinear specification provides a better fit.

Similarly, if L_1 and L_0 refer to the respective values of the maximized log likelihood function for the loglinear and linear specifications in Table 5, then $(L_1 - L_0)$ has the same functional form as in the preceding paragraph. For our sample, $(L_1 - L_0)$ equals 0.1186, so that again the loglinear specification is better.

Now say that L_1 is the maximized value of the log likelihood function for the loglinear specification of Table 5, where Kiefer's technique is used. Further, suppose that L_0 is the maximized value of the log likelihood function for the loglinear specification when Δ equals zero. Alternatively, the probability of being in a cooperative regime, λ, is one (or zero) for the entire sample. Then

$$L_1 - L_0 = \log\{(\|\hat{B}_1\| - \tfrac{1}{2}|\hat{\Sigma}_1|) - (\|\hat{B}_0\| - \tfrac{1}{2}|\hat{\Sigma}_0|)\}.$$

Under the null hypothesis that no regime change is observed, $2T(L_1 - L_0)$ has a chi-squared distribution with one degree of freedom. For the JEC sample, $2T(L_1 - L_0)$ is 554.1. Thus I can overwhelmingly reject the null hypothesis that no switch occurred, given the specifications adopted.

Among the models that have been estimated, the loglinear model presented in Table 5 that was estimated by Kiefer's method seems to have the most explanatory power. In particular, I can reject the alternative hypothesis that no switches in firm behavior took place, so that price and quantity changes cannot be solely attributed to exogenous changes in demand and structural conditions. The similarity of the estimated PN series and the PO series indicate that some price changes can be attributed to periods of noncooperative behavior, and that the incidence of alleged switches in behavior can't be explained by missing data problems.

6. Summary

The econometric evidence presented in the previous section indicates that reversions to noncooperative behavior did occur in the JEC, with a significant decrease in market price in these periods. The econometric results indicate that these episodes were concentrated in 1881, 1884 and 1885, which is consistent with the behavior of the JEC that was reported at that time. The question remaining, however, is what the cause of these reversions were.

Traditionally, breakdowns in cartel discipline have been attributed to demand slumps, both within the JEC as well as other cartels. What distinguishes the theoretical model of Green and Porter [5] from other theories of cartel stability is that reversionary episodes, or price wars, are caused by an unanticipated change in demand, in this case reflected by an unusually low market share for at least one firm, rather than a prolonged drop in total market demand. Trying to determine which model best describes the observed behavior of the JEC from 1880 to 1886 is not an easy task, but I can refer to two pieces of evidence which may support the Green and Porter paradigm. First, the reduced form estimates predict that price was lower and demand higher in noncooperative periods, ceteris paribus. Of course, this could merely reflect the fact that demand was quite elastic with respect to price changes, a fact at least partially refuted by the estimated price elasticity of demand. Secondly, one can look at total grain shipments from Chicago and see what fraction is accounted for by the JEC. Annual data showing the amount of grain shipped by lake steamers versus railroads is presented in Table 9. Of the years in the sample, 1880 is a boom year, which would account for the unusually high prices charged then. Of the remaining years, there isn't much annual variation in total shipments. Rather, the distinguishing feature of the "breakdown" years of 1881, 1884 and 1885 is the much higher market share captured by the JEC as a whole in the intermodel competition to ship wheat. This is an indication that JEC price wars were not concurrent with similar lake steamer price wars, and also that JEC price wars did not occur in years when total demand was unusually low. In fact, railway shipments were at their largest in 1881, 1884 and 1885. Thus while some observers have claimed that price wars will be triggered by the unexpected tapering off of demand, consistent with the paradigm of Green and Porter, the JEC seems to be a case where this was not necessarily true of periods in which demand was low per se. Further support of this contention is that the PO and PN series, as shown in Table 7, are not systematically related to the opening of the lake steamer shipping season. Finally, the fact that the frequency of noncooperative periods increased as the number of market participants increased is consistent with a story of dynamic cartel enforcement mechanisms, especially since the "no-exit" constraint faced by railroads deterred predatory reactions to entry.

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Table 1 : List of Variables[1]

GR	grain rate, in cents per 100 lbs.
TQG	total quantity of grain shipped, in tons.
LAKES	dummy variable; 1 if Great Lakes were open to navigation 0 otherwise
PO	cheating dummy variable; 1 if colluding reported by <u>Railway Review</u>, 0 otherwise.
PN	estimated cheating dummy variable.
DM1	1 from week 18 in 1880 to week 10 in 1883 0 othersise; reflecting entry by the Grand Trunk Railway
DM2	1 from week 11 to week 25 in 1883 0 otherwise; reflecting an addition to New York Central
DM3	1 from week 26 in 1883 to week 11 in 1886 0 otherwise; reflecting entry by the Chicago and Atlantic
DM4	1 from week 12 to week 16 in 1886 0 otherwise; reflecting acquisition of the Chicago and Atlantic

[1] The sample is from week 1 in 1880 to week 16 in 1886.

Table 2: Summary Statistics

Variable	Mean	Standard Deviation	Minimum Value	Maximum Value
GR	.2465	.06653	.125	.40
TQG	25384	11632	4810	76407
LAKES	.5732	.4954	0	1
PO	.6189	.4864	0	1

Table 3: Sample Correlation Matrix

	GR	TQG	LAKES
TQG	-.2543		
LAKES	-.3216	-.3173	
PO	.5776	-.3109	-.0299

Table 4: Two Stage Least Squares Results
(Employing PO)[1, 2]

	LINEAR		LOGLINEAR	
	Demand	Supply	Demand	Supply
C	46282	.2179	9.169	-3.964
	(11.87)	(5.19)	(49.7)	(-2.25)
LAKES	-13501		-.4367	
	(-4.57)		(-3.64)	
GR	-71499		-.7420	
	(-5.61)		(-6.15)	
DM1		-.0628		-.2014
		(-5.04)		(-3.69)
DM2		-.0540		-.1721
		(-2.96)		(-2.16)
DM3		-.0978		-.3220
		(-6.76)		(-5.04)
DM4		-.0681		-.2082
		(-1.88)		(-1.22)
PO		.8081		.4026
		(7.13)		(6.81)
TQG		$.2337 \times 10^{-5}$		.2506
		(1.79)		(1.47)
R^2	.322	.340	.312	.320
S	9787.33	.05536	.3975	.2430

[1] Monthly dummy variables are employed. To economize on space, their estimated coefficients are not reported.

[2] Estimated t-statistics are in parentheses.

Table 5: Maximum Likelihood Estimates (Yielding PN)[1]

	LINEAR		LOGLINEAR	
	Demand	Supply	Demand	Supply
C	49852	.2165	9.090	-2.438
	(15.2)	(10.0)	(61.0)	(-3.43)
LAKES	-13197		-.4305	
	(-4.45)		(-3.58)	
GR	-85130		-.7995	
	(-8.59)		(-8.81)	
DM1		-.0515		-.1648
		(-8.13)		(-6.76)
DM2		-.0628		-.2079
		(-6.75)		(-5.81)
DM3		-.0850		-.2856
		(-12.3)		(-10.5)
DM4		.0383		-.2909
		(1.67)		(-3.96)
PN		.1195		.5452
		(16.0)		(16.8)
TQG		$-.1363 \times 10^{-5}$		.0922
		(2.19)		(1.34)
R^2	.315	.827	.307	.863
S	9844.28	.02832	.3988	.1091

[1] PN is the estimated probability of collusion series, $(\hat{\omega}_1,\ldots,\hat{\omega}_T)$. The coefficient attributed to PN is the estimate of β_3.

Table 6: Price, Quantity, and Total Revenue for Different Values of LAKES and PN[1]

6.a PRICE

		LAKES 0	LAKES 1
PN	0	.1673	.1612
	1	.2780	.2679

6.b QUANTITY

		LAKES 0	LAKES 1
PN	0	38680	25904
	1	25775	17261

6.c TOTAL REVENUE[2]

		LAKES 0	LAKES 1
PN	0	32356	20878
	1	35872	23121

[1] Computed from the reduced form of the loglinear estimates of Table 5, with all other explanatory variables set at their sample means.

[2] Total Revenue = 5(Price x Quantity), to yield dollars per week.

Table 7: Predicted Incidence of Noncooperative Episodes

	Beginning of Episode Year	Week	Length of Episode (in Weeks)
PO Series	1881	10	8
	1881	24	33
	1883	8	1
	1883	17	2
	1883	36	2
	1884	7	19
	1884	31	10
	1885	3	3
	1885	11	37
	1886	3	9
PN Series[1]	1881	27	19.4
	1881	49	15.0
	1884	3	1.0
	1884	13	13.2
	1884	36	6.0
	1885	2	4.4
	1885	12	5.0
	1885	20	11.0
	1885	34	8.0
	1885	49	4.4

1. This is taken from the estimated PN series in the loglinear system.

Table 8: Index of Cartel Non-Adherence[1]

Year	MacAvoy[2]	Ulen[3]	Estimated[4]
1880	26	0.00	0.00
1881	14	0.67	0.45
1882	18	0.06	0.21
1883	6	0.10	0.00
1884	16	0.58	0.39
1885	10	0.77	0.62
1886[5]	15	0.50	0.03

1. Columns 1 and 2 are taken from Ulen [15], Table 17, page 336.

2. The total number of months for which, for all cartel members, the difference between the actual market share and "trend" share of tonnage was greater than the standard error from the "trend" share regression of each member road. The greater this number of months, the less stable the cartel is likely to be.

3. If period t is in year i, then say $t \in T_i$. Then this index is $\Sigma(1 - PO(t))/\Sigma t$, where the summation is over t in T_i.

4. This index is $\Sigma(1 - \hat{w}_t)/\Sigma t$, summing over t in T_i for each year i.

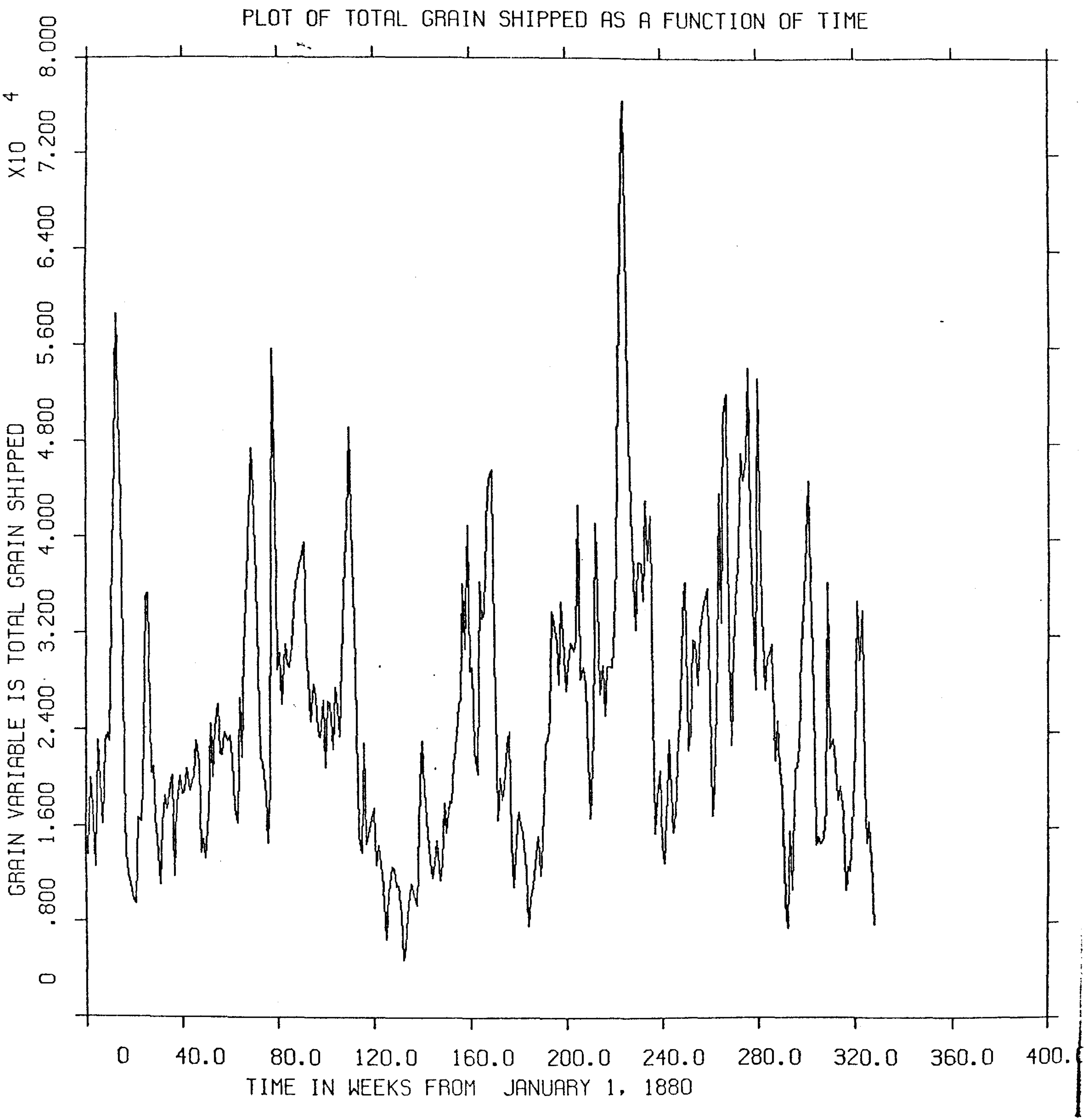
5. PO and PN only exist for the first 16 weeks, so $T_i = 16$.

Table 9: Annual Eastbound Shipments of Wheat From Chicago by Lake and Rail

Year	Lake		Rail	
	Total[1]	Percentage	Total[1]	Percentage
1880	16.69	77.9	4.728	22.1
1881	7.688	50.0	7.680	50.0
1882	14.94	86.2	2.389	13.8
1883	7.067	73.2	2.590	26.8
1884	11.52	66.0	5.928	34.0
1885	5.436	51.5	5.116	48.5
1886	10.51	82.6	2.209	17.4

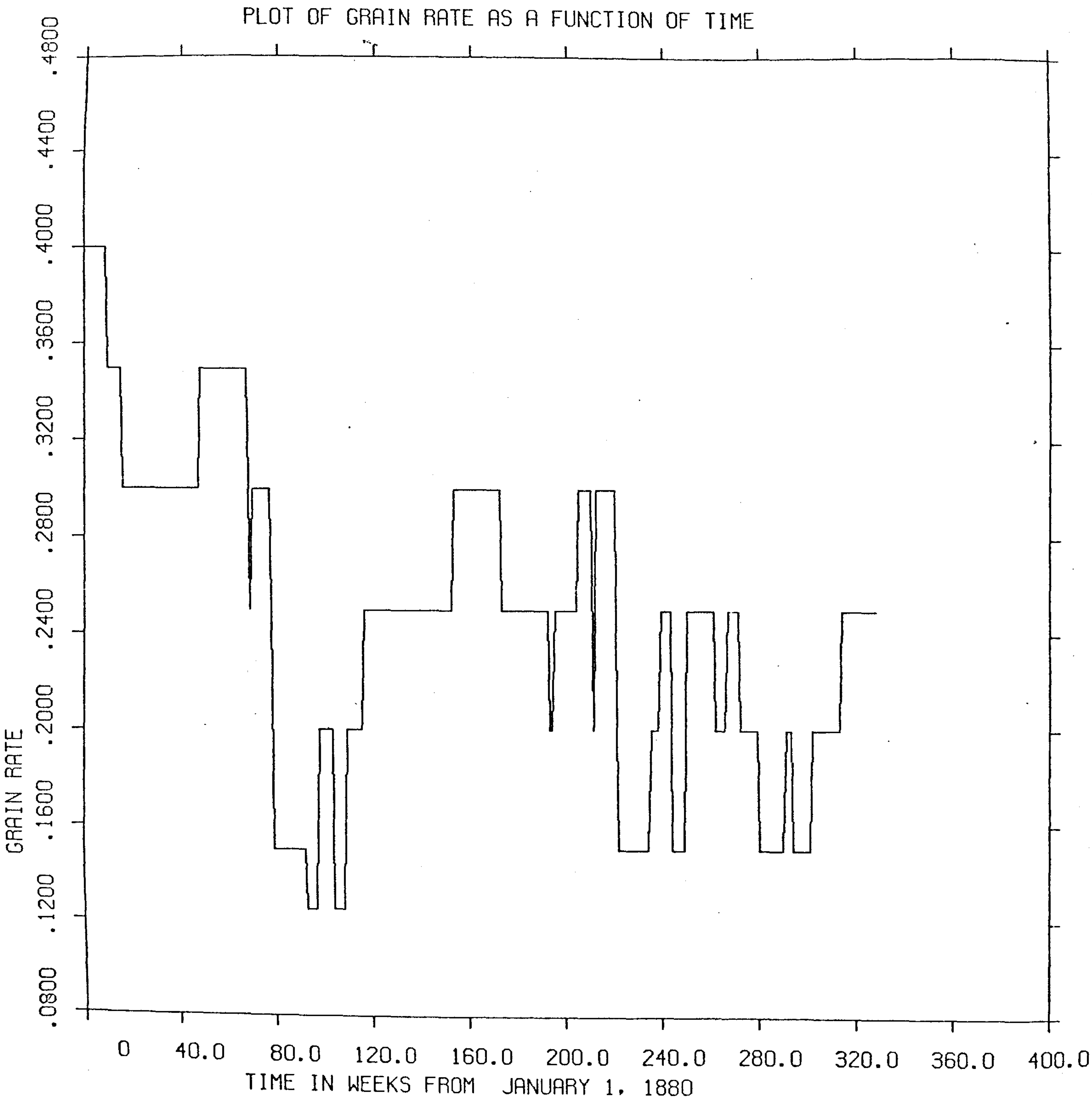
[1] in millions of bushels.

FIGURE 1



PLOT OF TOTAL GRAIN SHIPPED AS A FUNCTION OF TIME
GRAIN VARIABLE IS TOTAL GRAIN SHIPPED
X10 4
0
.800
1.600
2.400
3.200
4.000
4.800
5.600
6.400
7.200
8.000
0
40.0
80.0
120.0
160.0
200.0
240.0
280.0
320.0
360.0
400.0
TIME IN WEEKS FROM JANUARY 1, 1880

FIGURE 2


PLOT OF GRAIN RATE AS A FUNCTION OF TIME
GRAIN RATE
.0800
.1200
.1600
.2000
.2400
.2800
.3200
.3600
.4000
.4400
.4800
0
40.0
80.0
120.0
160.0
200.0
240.0
280.0
320.0
360.0
400.0
TIME IN WEEKS FROM JANUARY 1, 1880