Price Wars and the Stability of Collusion:
A Study of the Pre-World War I Bromine Industry

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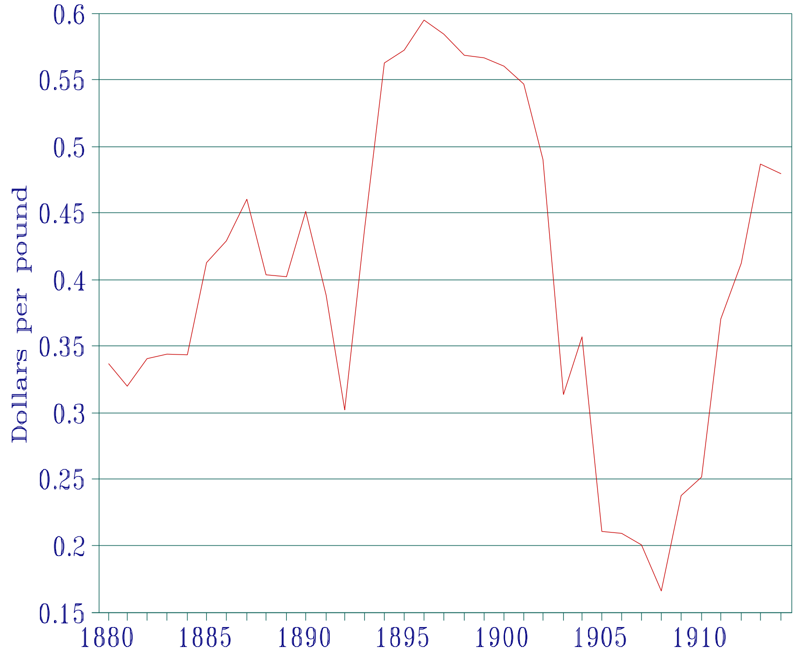
Patterns in the Real Price of Bromine and Potassium Bromide, 1860-1914

The real price of bromine fell over 90% between 1860 and 1880. It fell another 18.4% from 1880 till the founding of the National Bromine Company in 1885. The establishment of the NBC reversed this trend; the real price of bromine increased 15.6% during 1885. The average real price of potassium bromide was consistently higher during cooperative than non-cooperative periods, both before and after Dow's entry. The average real price of potassium bromide was 21% higher during the NBC pool than during the previous five years. At the expiration of the NBC, the price fell over 40%. The average price of potassium bromide during the Shields pool was 64% higher than during the period between the NBC and Shields pools. When the Shields contracts expired in 1902, the price dropped 32% in one month. After 1902, real potassium bromide prices were, on average, 74% higher during cooperative periods than non-cooperative periods.

Nominal bromine and potassium bromide prices from the Oil, Paint, and Drug Reporter have been deflated using wholesale chemical price indices from the United States Historical Statistics (series E49 and E60), pp. 200-201.



Real Price of Potassium Bromide, 1880-1914

Tests of the Markov Structure of the Timing of Price Wars in the Bromide Industry

In the Abreu, Pearce, and Stacchetti [1986] model, the industry remains in a collusive (price war) state until a public variable falls into a range indicating, with sufficient likelihood, that a firm has not played (has played) its prescribed strategy, generating a first-order Markov process. If the state of the industry follows a first-order Markov process, then the probability of being in a particular state in period *t* will depend only on the state of the industry in period $t-1$. More generally, if it follows an *n*-order process, then the transition probabilities will depend on the history $t-1, ..., t-n$.

Berry and Briggs [1988] suggest a procedure for testing whether an industry which is known to have switched from collusive to price-war phases was implementing an APS-type mechanism. The test asks whether a variable indicating the state of the industry follows a first-order Markov process, as predicted by APS. If that were the case, it would imply that industry participants based their actions in the current period only on what occurred in the immediately prior period. In other possible, more complex punishment strategies, including those suggested by Green-Porter, one's action today might depend as well on the state of the industry in more distant periods. Thus the test asks whether the transition probability, $P(I_t = 1)$, is the same following histories which are the same for the immediately prior period, $t-1$, but differ for more distant periods. If the estimates of transition probabilities following similar one-period, but different *n*-period histories ($P(I_t = 1)|H_i$, $P(I_t = 1)|H_j$, where $H_{it-1} = H_{jt-1}$, $H_{it-n} \neq H_{jt-n}$) are different, then the test rejects the null of a first-order Markov process for the alternative of an *n*-order process. Berry and Briggs use this procedure to analyze data from the Joint Executive Committee and conclude that it was more likely to have been implementing an APS type agreement than a Green-Porter one (as concluded in Porter [1983]).

I replicate their test for the bromine industry for the period 1885-1914. I construct two series, both of which measure the state of collusion in the industry. The first series has 1539 observations of periods of one week length. This was the frequency of public price announcements. The second series has 354 observations of one month periods, more closely corresponding to the time it took to arrive at a coordinated response, given the geographical distance separating firms. Each series $\{I_t\}$ is defined such that $I_t = 1$ if firms in the industry are colluding that period, $I_t = 0$ otherwise. The data are based on published reports of collusive activity in the *Oil, Paint and Drug Reporter* and the private papers of the Dow Chemical Company and other firms in the industry. Price data were not used to construct these series. I then ask whether the estimates of the transition probabilities implied by these data, following similar one-period histories, are equivalent following different *n*-period histories. The results of these tests are presented in Table I.

The bromine industry results resemble those obtained by Berry and Briggs. Where it is possible to test formally the hypothesis of a first-order Markov structure, it is impossible to reject the null that more distant periods were irrelevant: the probabilities of colluding in period *t*, following two different *n*-period histories which were in the same state in period *t-1*, were statistically indistinguishable. This is the case whether one considers relatively short periods (one week) and short histories (two or three weeks) or whether one uses longer periods (one month) and considers the possibility that more distant periods - up to five months of prior history - might have an impact on transition probabilities in the current period.

Where formal hypothesis testing was not possible, because the variance of an estimate of the probability of switching from one state to another is zero, we can calculate the probability that we would observe a zero variance, given a first order Markov process and the estimated transition probabilities. In no case is the probability of observing a zero variance less than 68%. Thus the general pattern and timing of price wars is not inconsistent with that predicted by APS.

Intuitively, the zero variance estimates arise because there is little switching back and forth between price wars and cooperation. For example, every single history ($I_{t-2} = 0, I_{t-1} = 1$) in the sample is followed by $I_t = 1$. (See Tables II and III.) That is, there is no case in which a price war was followed by a period of cooperation and then an immediate resumption of the price war. Thus the estimate of the $P(I_t = 1| H = (I_{t-2} = 0, I_{t-1} = 1))$ is equal to one, and the estimate of the dispersion around that estimate is zero. Similarly, there are no cases of cooperation lasting exactly two periods, or of price wars lasting exactly two periods. Once the industry entered a cooperative or price war state, that state almost always endured for some length of time. The probability of the industry transitioning out of the state it was in was low, especially if the industry was currently cooperating. If monitoring problems had been more severe, and the probability of beginning a price war greater, we would have more observations of transitions between price wars and collusion, and have fewer zero variance estimates of the probability of making that transition. But collusion in the bromine industry was more successful than that. While we have over 15,000 weekly observations of the state of the industry, with only six price wars, we have only twelve observations of transitions from one state to another.

Because the Berry-Briggs test requires that the Markov-one hypothesis be compared with an alternative of specific length, it seems possible that the "cannot reject APS" result that this test gives could be the result of specifying an incorrect alternative. Thus we turn to another, more general, test. In this case, we construct a theoretical distribution of the duration of price wars (using the estimated transition probabilities presented in Tables II and III) under the Markov-one hypothesis, and then compare that theoretical distribution with the distribution of the observed price war lengths. The Kolmogorov-Smirnov goodness of fit test asks whether there is a statistically significant difference between the two distributions. In other words, given that there were six price wars during the period studied, is the length of those six price wars consistent with their having been generated by a Markov-one process? Using both monthly and weekly data, we again find that we cannot reject the null hypothesis of a first order Markov process (Table I).

Thus the statistical tests allow us to say that the observed transitions between collusion and price wars are consistent with APS strategies generating a first-order Markov process. However, because the alternative hypothesis specified is always a Markov process of some longer length, the findings are also consistent with some other underlying (non-Markov) process in which the probability of commencing a price war is lower when the industry has just emerged from a price war. In that case, the transition probabilities would not be constant, but rather a changing function of other observables, as with the complex strategies that the APS strategy avoids. Such a history-dependent strategies might also explain the very small numbers of switches back and forth between states which generated the zero-variance estimates in the first set of tests. For example, it might be that the probability of beginning a price war is lower in the periods immediately following the end of a price war, rather than constant over all collusive periods. Thus while the statistical evidence is not inconsistent with the one empirically testable conclusion of APS, it is also not inconsistent with much more complex, history dependent strategies.

TABLE I

TESTS OF MARKOV STRUCTURE OF PRICE WAR OCCURRENCE

Test	*Null*	*Alternative*	*Test Statistic*	*Significance level*	*Degrees of Freedom*
			Weekly Series		
Berry-Briggs	M(0)	M(1)	18139.06	.999	1
Kolmogorov-Smirnov	M(1)	Not M(1)	0.35	.65	

Note: Tests of M(2), M(3), M(4), and M(5) versus M(1)cannot be performed with weekly data because $Var(P(I_t=1|H_i)) = 0$ for most two, three, four, and five-period histories. See Table II.

PROBABILITY OF ZERO VARIANCE UNDER A MARKOV ONE STRUCTURE

If $P(I_t=1|I_{t-1}=0) = .018$, then P(6 zeroes in a row) = .897.

If $P(I_t=1|I_{t-1}=1) = .994$, then P(6 ones in a row) = .965.

If $P(I_t=1|I_{t-1}=0) = .018$, then P(12 zeroes in a row) = .804.

If $P(I_t=1|I_{t-1}=1) = .994$, then P(12 ones in a row) = .930.

			Monthly Series		
Berry-Briggs	M(0)	M(1)	844.04	.999	1
Berry-Briggs	M(1)	M(2)	0.41	.477	1
Berry-Briggs	M(1)	M(3)	0.36	.450	1
Berry-Briggs	M(1)	M(4)	0.96	.381	2
Berry-Briggs	M(1)	M(5)	1.36	.493	2
Kolmogorov-Smirnov	M(1)	Not M(1)	0.34	.61	

PROBABILITY OF ZERO VARIANCE UNDER A MARKOV ONE STRUCTURE

If the $P(I_t=1|I_{t-1}=1) = .974$, then P(6 ones in a row) = .854.

If $P(I_t=1|I_{t-1}=0) = .074$, then P(5 zeroes in a row) = .681.

If the $P(I_t=1|I_{t-1}=1) = .973$, then P(12 ones in a row) = .720.

Table II

ESTIMATED TRANSITION PROBABILITIES FOR WEEKLY DATA SERIES

History (t-3,t-2,t-1)	*$P(I_t=1\|H_i)$*	*Var(Prob)*	*N*
	One Period Histories		
(*,*,0)	0.017	0.0167	394
(*,*,1)	0.995	0.0050	1195
	Two Period Histories		
(*,0,0)	0.018	0.0177	338
(*,1,0)	0.000	0.0000	6
(*,0,1)	1.000	0.0000	6
(*,1,1)	0.994	0.0060	1188
	Three Period Histories		
(0,0,0)	0.018	0.0177	332
(1,0,0)	0.000	0.0000	6
(1,1,0)	0.000	0.0000	6
(1,1,1)	0.994	0.0060	1181
(0,1,1)	1.000	0.0000	6
(0,0,1)	1.000	0.0000	6

Table III

ESTIMATED TRANSITION PROBABILITIES FOR MONTHLY SERIES

History (t-3,t-2,t-1)	$P(I_t=1\|H_i)$	*Var(Prob)*	*N*
	One Period Histories		
(*,*,0)	0.076	0.070	79
(*,*,1)	0.978	0.022	275
	Two Period Histories		
(*,0,0)	0.068	0.063	73
(*,1,0)	0.167	0.139	6
(*,0,1)	1.000	0.000	6
(*,1,1)	0.974	0.025	268
	Three Period Histories		
(0,0,0)	0.074	0.069	68
(1,0,0)	0.000	0.000	5
(1,1,0)	0.167	0.139	6
(1,1,1)	0.973	0.026	261
(0,1,1)	1.000	0.000	6
(0,0,1)	1.000	0.000	5
(1,0,1)	1.000	0.000	1

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