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Price Analysis of the Pinto and Great Northern Dry Edible Bean Market

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Price Analysis of the Pinto and Great Northern Dry Edible Bean Market

INTRODUCTION

The US dry edible bean industry has unique structural features that may impair the market's ability to allocate and signal. Edible beans are grown in geographically separated production regions; Southern Idaho, northeast Colorado, eastern North Dakota, Michigan, and western Nebraska-eastern Wyoming. While most edible bean varieties are generally production substitutes, due to specialization in processing and variety breeding programs, each production area is dominated by one or two varieties. In each production area there are a large number of farmers selling to a highly concentrated bean processor segment. Small processors or buyers and the even smaller grower cooperatives are confined to single production regions, while a few large processors have multi-region operations. Bean processors clean, package, store, and transact sales i.e. serve as the "middleman" between farmer and export market dealers.

The majority of edible beans are exported internationally into a volatile market. The major export markets in the Middle East, Mexico and Central America, and Africa are politically and economically unstable which translates into price uncertainty. Iraq is a major buyer of Great Northern beans. Mexico buys only Pinto beans. At the farm level, bean varieties are production substitutes. However at the market level little variety substitution is evidence -- each export market is dominated by a specific variety. Neither price nor volume of beans destined to these export markets is publicly available information.

There is no futures market for edible beans. In the absence of clear and timely national prices established by a futures market, USDA reports the prices of transactions for the dominant varieties in each major production region. Price information gaps occur when the major markets lack transactions in a certain time period. Farmers have limited information about the historical pricing behavior. A significant proportion of beans are contracted prior to harvest to a specific local processor and contract prices are not reported. Farmers thus have difficulty to determine current and future prices for different bean varieties to make production and marketing decisions. Some of the uncertainty over the price of a particular variety in a specific market area could be reduced, if there was information on how edible bean prices compare across production areas and among different varieties.

The objective of this bulletin is to analyze edible bean prices to 1) determine the relationship of prices in different production regions, 2) determine the relationship of prices for different bean varieties, and 3) determine the relationship between grower and dealer prices. In particular, the bulletin will focus more on the theoretical and statistical techniques to analyze these problems and relationship in the dry bean market.

Data

In this report, monthly grower and dealer prices for two major varieties: Pinto beans and Great Northern beans are investigated. USDA Livestock and Sees Division reports weekly grower prices and dealer prices for Pinto and Great Northern beans in different markets. Those weekly prices are averaged to calculate the monthly prices. Prices for the Pinto variety were collected from four major production regions for that variety: Colorado, North Dakota, Idaho, and combined region of Western Nebraska-Eastern Wyoming. Prices for the Great Northern variety were collected for two production regions for this variety: Idaho and the combined region of Western Nebraska-Eastern Wyoming. The price series covers the period from September 1983 to August 1996, totally 153 months.

Hypotheses

Several hypothesis are set forth and empirically tested with the above described data. The first hypothesis is that edible bean prices for the same variety will be cointegrated across the distinct production regions. A second and related hypothesis is that pinto and Great Northern prices will be cointegrated. Another hypothesis is that the dominant production region for each variety will be the price leader for that variety. Specifically, Colorado will be the price leader for pinto beans and western Nebraska-eastern Wyoming will be the price leader for Great Northern beans. The final hypothesis relates to the margin between grower and dealer prices. We hypothesize that the margin will be proportional to the variance in dealer prices.

This report is divided into major sections that outline the theory behind and methods used to test each hypothesis. The empirical results for each hypothesis are contained within each of these sections. A final conclusion section summarizes the findings and states the implications. However, before the first hypothesis on cointegration analysis can be tested, the stationarity of a price series must be determined.

PRICE STATIONARITY

Theory

With dry bean price data being recorded as a time-series and the first analysis undertaken of time series is the stationarity of the series. A time series is stationary if its mean, variance, and autocovariances are independent of time (Rao, 1994). Suppose y_t is a time series (or stochastic process) that is defined for $t = 1,2,...$ and for $t = 0,-1,-2,...$. Formally y_t is covariance (weakly) stationary if the following conditions are satisfied (Rao, 1994):

$$E(y_t) = \mu \tag{1}$$

$$E[\,(y_t - \mu)^2\,] = var(y_t) = x(0) \tag{2}$$

$$E[(y_t - \mu)(y_{t-\iota} - \mu)] = cov(y_t,\ y_{t-\iota}) = x(\iota), \qquad \iota = 1,2,3,\cdots \tag{3}$$

Equation (1) requires the process to have a constant mean with a value of μ. Equation (2) requires the process to have a constant variance $x(0)$ which is invariant of time. Finally equation (3) requires the covariance between any two values from the series (an autocovariance) depends only on the time interval between those two values (ι) and not on the point of time (t).

Define the autocovariance function in (3) as:

$$corr(y_t\,,\ y_{t-\iota}) = \frac{cov(y_t\,,\ y_{t-\iota})}{\sqrt{var(y_t)var(y_{t-\iota})}} = \frac{x(\iota)}{x(0)} \tag{4}$$

Consider an example of AR(1) process defined by

$$y_t = \rho\ y_{t-1} + e_t, \quad t = \cdots,-1,0,1,\cdots \tag{5}$$

where e_t is assumed to define a sequence of independently and identically distributed (*IID*) random variables with expected value zero and variance σ^2. The process in (5) is stationary when ρ is less than one in absolute value, i.e. $-1 < \rho < 1$ (Rao, 1994). This can be proved by introducing the lag operator, L, where $Ly_t = y_{t-1}$ and $L^2y_t = L(Ly_t) = Ly_{t-1} = y_{t-2}, \ldots$ and so on. Equation 5 can thus be rewritten as:

$$y_t - \rho\, y_{t-1} = y_t - \rho\, L\, y_t = (1 - \rho\, L)\, y_t = e_t \tag{6}$$

so that $y_t = (1 - \rho L)^{-1}\, e_t$. If we let $p(L) = 1\text{-}\rho L$, we can re-write (6) as

$$p(L)\, y_t = e_t \tag{7}$$

The root of $p(L)$ is given by $L = (1/\rho\,)$ so that the requirement for ρ has absolute value less than one is equivalent to requiring that the root of $p(L)$ is greater than one in absolute value. Furthermore, $p(L)$ has a unit root, i.e. the AR(1) has a unit root, if and only if $\rho = 1$. In this case the stationarity condition is not satisfied, and the AR(1) process with a unit root is non-stationary. This implication can be explored by contrasting the unit root case with the stationary case.

Assume the process starts at $t = 0$, thus

$$y_t = \rho\, y_{t-1} + e_t\,, \qquad t = 1,2,3,\cdots \tag{8}$$

where y_0 is a fixed initial value for the process. By repeating backwards substitution in (9) one can get

$$y_t = \rho^t\, y_0 + e_t + \rho\, e_{t-1} + \rho^2\, e_{t-2} + \cdots + \rho^{t-1}\, e_1 \tag{9}$$

Now assume that $\rho = 1$, then

$$y_t = y_0 + e_t + e_{t-1} + e_{t-2} + \cdots + e_1 = y_{t-1} + e_t \tag{10}$$

and:

$$\begin{aligned} E(y_t) &= y_0 \\ var(y_t) &= t\,\sigma^2 \\ corr(y_t\,,\,y_{t-l}) &= \sqrt{\frac{t-l}{t}} \end{aligned} \tag{11}$$

In the case where $|\rho| < 1$:

$$\begin{aligned} E(y_t) &= \rho^t\, y_0 \\ var(y_t) &= \sigma^2\,[\frac{1-(\rho^2)^t}{1-(\rho^2)}] \\ corr(y_t\,,\,y_{t-l}) &= \rho^l \sqrt{\frac{1-(\rho^2)^{t-l}}{1-(\rho^2)^t}} \qquad l = 1,2,3,\ldots \end{aligned} \tag{12}$$

By looking at (12), it seems that when $|\rho| < 1$ the series does not satisfy the stationary conditions since, for example, var(y_t) depends on time. However the limits for (12) can be taken and the case can be called stationary,

$$\begin{aligned} \lim_{t\to\infty} E(y_t) &= 0 \\ \lim_{t\to\infty} var(y_t) &= \frac{\sigma}{1-\rho^2} \\ \lim_{t\to\infty} corr(y_t\,,\,y_{t-l}) &= \rho^l \end{aligned} \tag{13}$$

In general, assuming $|\rho| < 1$ the processes in (5) and (9) are equivalent when $t\to\infty$. Thus if t = 1,2,3,...,T defines the values of y_t that are observed, the least square estimator of ρ is

$$\hat{\rho} = \frac{\sum_{t=2}^{T} y_t\, y_{t-1}}{\sum_{t=2}^{T} y_{t-1}^{\;2}} \tag{14}$$

Then

$$\sqrt{T}\,(\hat{\rho} - \rho) \rightarrow N(0\ ,\ 1-\rho^2) \tag{15}$$

whether the AR(1) process starts at $t = 0$ or in the infinite past.

In the case where $\rho = 1$ the variance of y_t increases without bound as $t \rightarrow \infty$, and

$$\lim_{t \rightarrow \infty} corr(y_t\ ,\ y_{t-l}) = 1, \qquad l = 1,2,3,\cdots \tag{16}$$

which contrasts with the result in (13) where corr(y_t, y_{t-l}) fades away as ρ increases. For a stationary series the estimated autocorrelations should fade away rapidly as ρ increases whereas for a non-stationary series they should not tend to fade away. Since the following is true in the unit root case,

$$\frac{\partial y_t}{\partial e_{t-s}} = 1, \qquad s = 1,2,3,\cdots \tag{17}$$

and the following is true in the stationary case,

$$\frac{\partial y_t}{\partial e_{t-s}} = \rho^s, \qquad s = 1,2,3,\cdots \tag{18}$$

then, a "shock" or "innovation" has a sustained effect in the unit root case, while a similar shock has a diminished effect over time in the stationary case.

Methods

The Dickey-Fuller (DF) Unit Root Test for a univariate time series is applied to the price series. The null hypothesis is the existence of a unit root for each series. While the alternative hypothesis states the series is stationary. SAS/ETS has a macro procedure for DF tests, and details about the testing procedures can be found in Hamilton (1994) and the SAS/ETS User's Guide.

Empirical Results

Figure 1 to Figure 6 show the historical data of dealer prices and grower prices for Pinto and Great Northern Beans in selected markets. Table 1 shows the p-value for the unit root test on dealer price and grower price for each variety in each different market.

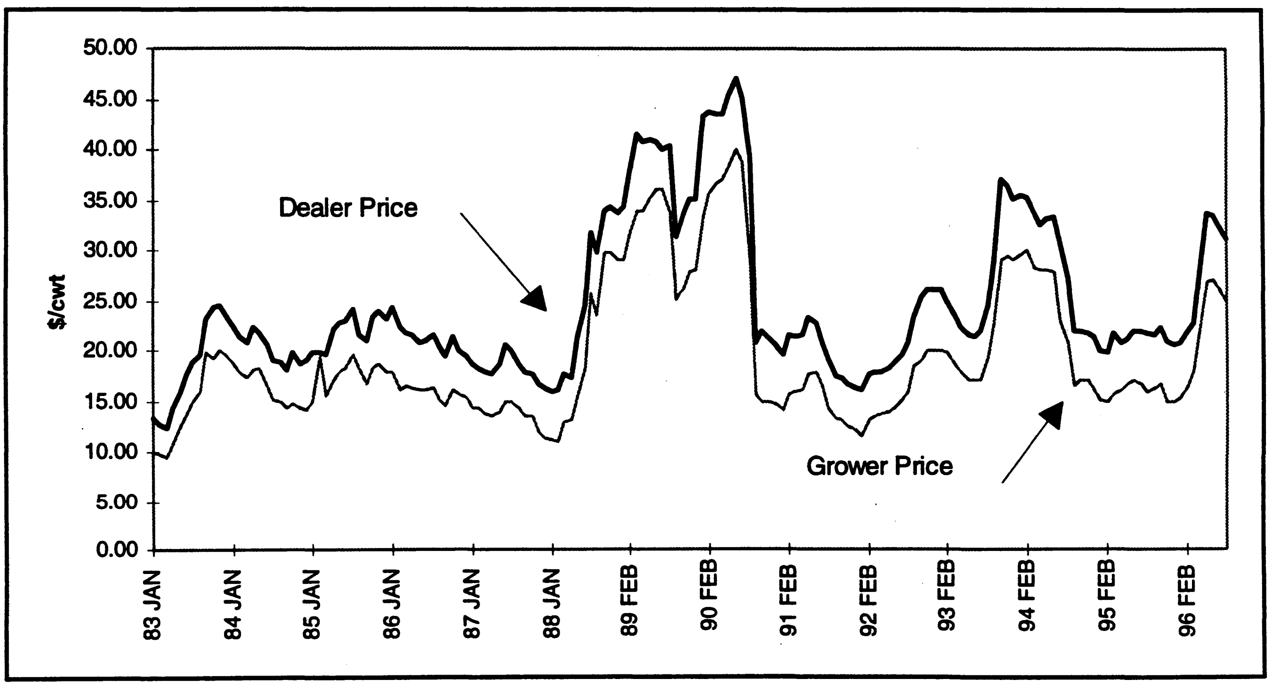

Figure 1. Historical Dealer and Grower Prices for Pinto Beans in Colorado, 1983-1996.

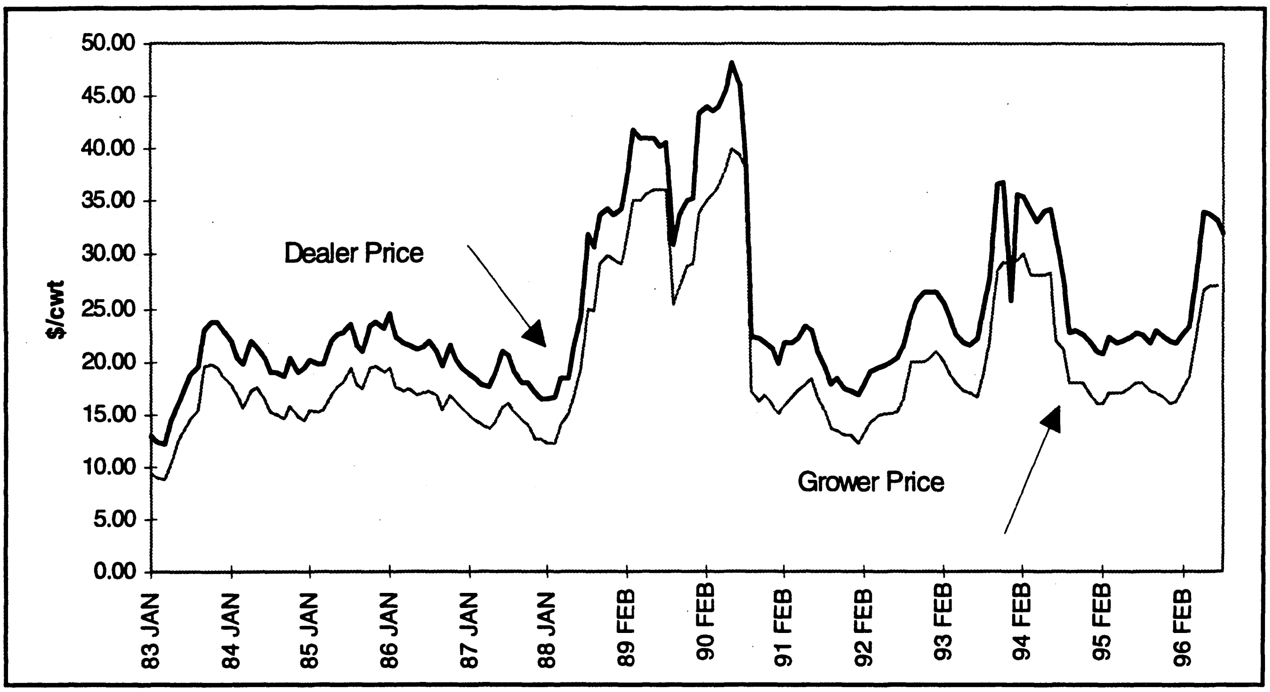

Figure 2. Historical Dealer and Grower Prices for Pinto Beans in Idaho, 1983-1996.

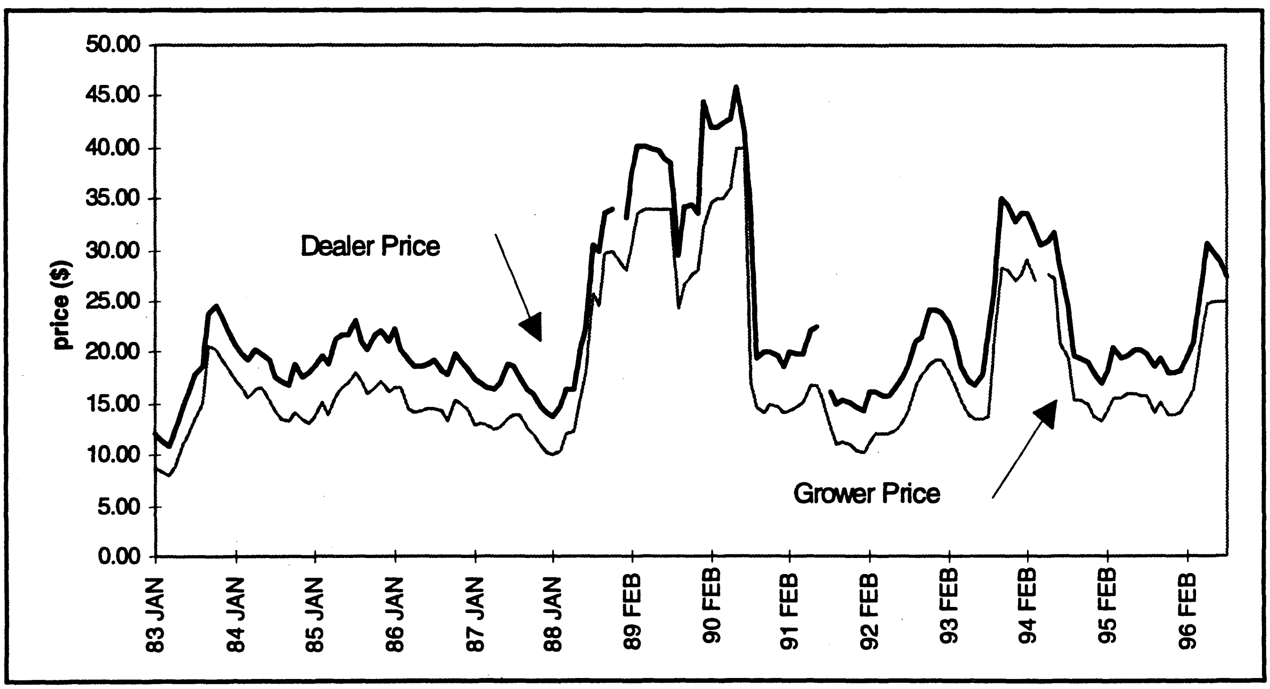

Figure 3. Historical Dealer and Grower Prices for Pinto Beans in North Dakota, 1983-1996.

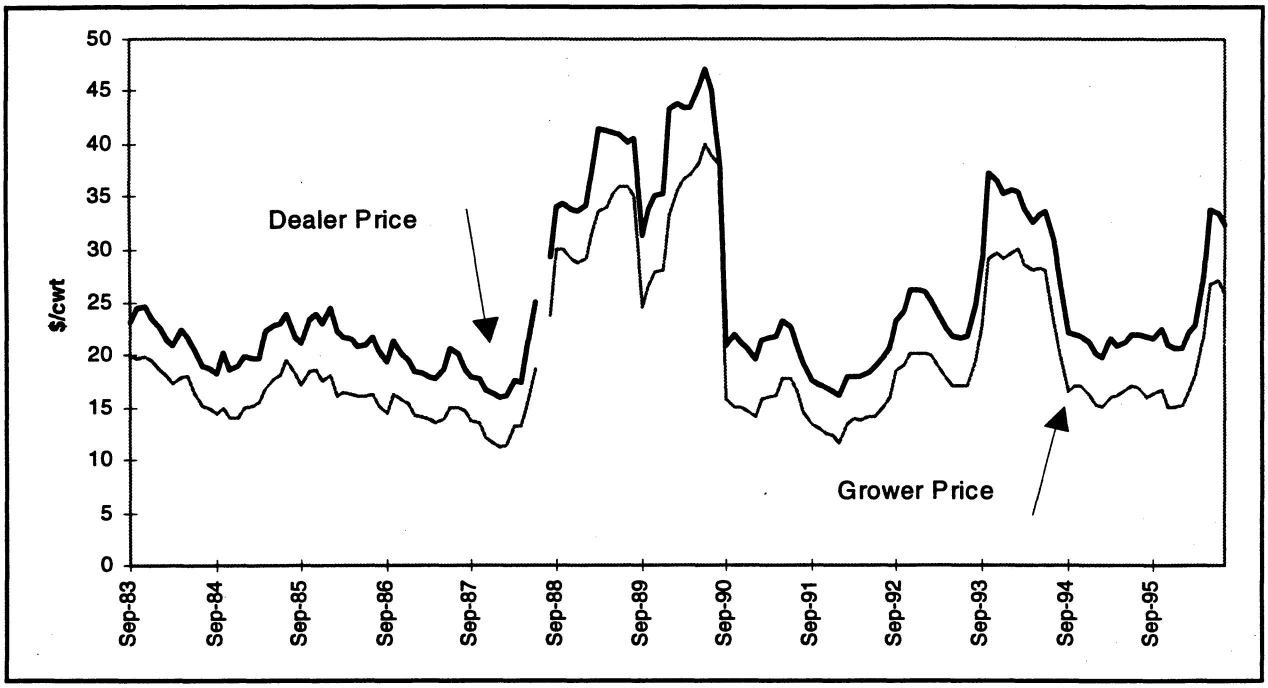

Figure 4. Historical Dealer and Grower Prices for Pinto Beans in W. Nebraska-E. Wyoming, 1983-1996.

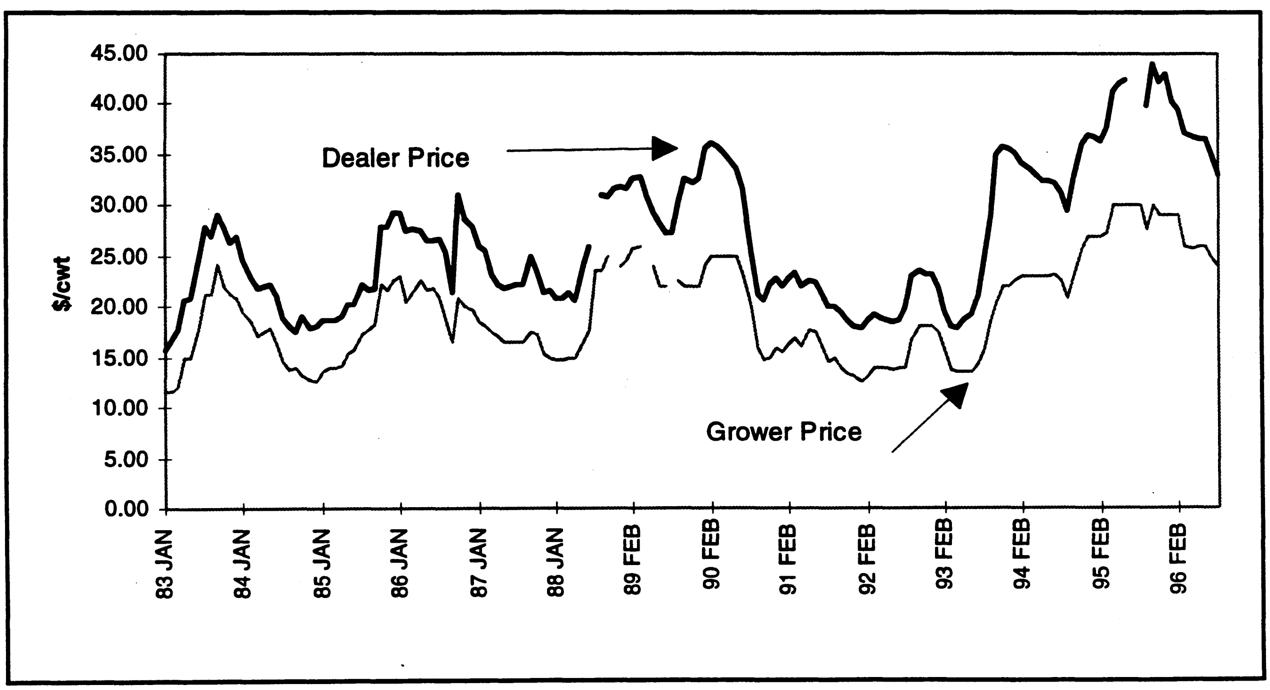

Figure 5. Historical Dealer and Grower Prices for Great Northern Beans in Idaho, 1983-1996.

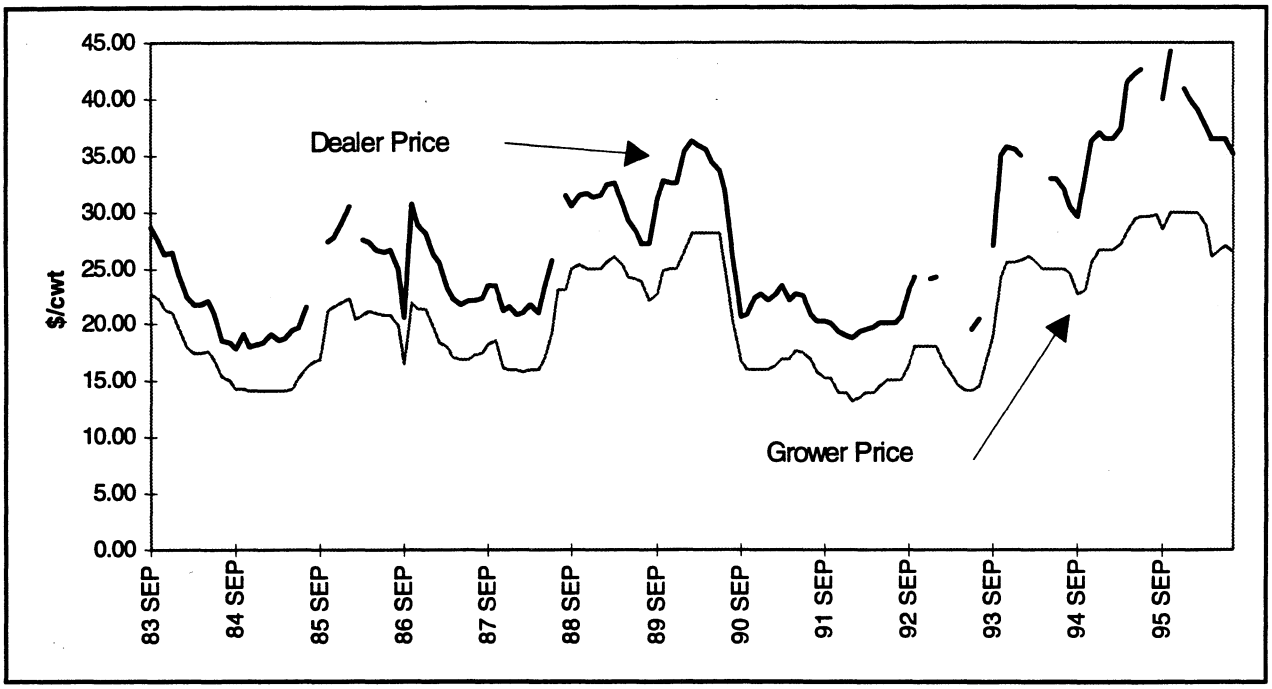

Figure 6. Historical Dealer and Grower Prices for Great Northern Beans in W. Nebraska-E. Wyoming, 1983-1996.

Table 1. P-value of Unit Root Test on Dealer and Grower Prices for Each Variety in Different Markets.

Variety	Market	Dealer Price	Grower Price
Pinto	Colorado	0.053	0.039
	Idaho	0.067	0.050
	North Dakota	0.043	0.055
	Neb. - Wyo	0.054	0.040
Great Northern	Neb. - Wyo.	0.233	0.182
	Idaho	0.248	0.160

Looking at Figures 1 through Figure 6, it seems that bean price series are not stationary. Some of them seem to have a slightly increasing trend. Unit root tests in Table 1 corroborate this conclusion. All p-values are significant at the 0.01 level. The null hypothesis of an existence of a unit root for each price series can not be rejected. In other words, none of the dealer price or grower price series are stationary, and none of them have a stable mean value over time. Later in the section on price leadership among dealer prices for each bean variety, the data is transformed so that all the dealer price series become stationary series. Details on the data transformation are given in that section.

COINTEGRATION ANALYSIS

It was shown in the previous section that dealer prices and grower prices are not stationary for Pinto and Great Northern beans. Each price series varies over time, and did not have a constant mean or a constant variance. The next question to answer is: what kind of the relationship exists between dealer prices or grower prices among selected markets? We hypothesized that prices across production regions and between the two varieties would be cointegrated. In the economic sense, cointegration implies that several nonstationary economic variables tend to move together in the long run, due to common forces behind those variables.

Theory

Granger (1981, 1991), Granger and Weiss (1983), and Engle and Granger (1987) have shown that, even though a given set of series may be non-stationary, there may exist various linear combinations of the individual series that are stationary. The desire to estimate models that combine both short-run and long-run properties and that at the same time maintain stationarity in all of the variables, has prompted a reconsideration of the problem of regression using variables measured in their levels. Cointegration is a statistical framework to test for long-run or steady-state equilibrium relationships among several non-stationary series.

The formal definition of cointegration of two variables developed by Engle and Granger

(1987) is as follows:

Define: Time series x_{1t} and x_{2t} are said to be cointegrated of order d, b, where d≥b≥0 written as x_{1t}, x_{2t} ~ $CI(d,b)$ if

1. Both series are integrated of order d,
2. There exists a linear combination of these variables, say $\alpha_1 \times x_{1t} + \alpha_2 \times x_{2t}$, which is integrated of order (d-b)

The vector [α_1, α_2] is called a cointegrating vector. If there is a long-run relationship between two (or more) nonstationary variables (all integrated of the same order), the idea is that deviations from this long-run path are stationary if the variables are to be cointergated.

Cointegration Tests on Spatial Price Relationship

Spatial market integration necessarily imply a unique long-run equilibrium relationship in which deviations from regional price parity are forced to zero (Goodman and Schroeder, 1991). Pinto and Great Northern beans are nearly perfect substitutes for the farmers in regional production. The two varieties can be planted with identical management practices and schedules. There is little costs to farmers when shifting between Pinto and Great Northern beans. Differences in Pinto or Great Northern prices in different markets should be exactly equal to transportation costs plus other constant costs, if these prices are affected by the same factors of changes in demand or changes in supply. As for the two varieties in the same market, growers might expect two price series to follow each other closely since they act like production substitutes. Dealers and growers in different markets would also expect their prices to move closely together in different markets, given the assumed perfect competition situation. The specific hypotheses to be tested in this section are:

(i) if dealers act as perfect competitors, dealer prices in selected markets should move closely together; i.e., dealer prices should be cointegrated.

(ii) if growers act as "perfect price takers" under perfect competition, grower prices should move closely together in selected markets; i.e., grower prices should be cointegrated.

(iii) if Pinto and Great Northern beans are perfect substitutes for each other, then Pinto grower prices should be cointegrated with Great Northern grower prices.

Procedures for Cointegration Tests

Consider two price series in the following regression:

$$p_t^1 - \alpha - \beta\, p_t^2 = u_t \tag{19}$$

where p_t^1 and p_t^2 represent Pinto prices in two markets, for example. Existence of perfectly spatially integrated markets (where price changes in one market are fully reflected by equilibrating changes in alternative market) necessarily requires that the estimated parameter of the cointegrating regression, β , have a value of one. However, because the price series p_t^1 and p_t^2 are nonstationary in a cointegrated system, conventional *t*-test cannot be used to provide reliable hypothesis tests regarding the value of β.

We adapt four testing procedures suggested by Engle and Granger (1987) for cointegration. They also provided critical values for a sample of 100 observations based on the results of Monte Carlo simulations for each proposed test statistics. Null hypothesis for each test is "no cointegration". Rejection of null hypothesis affirms integrated prices in regional markets with this case study.

1. The Co-integrating Regression Durbin Watson:

$$y_t = \hat{\alpha}\, x_t + c + \hat{e}_t \tag{20}$$

Test Statistic $= DW =$

$$\frac{\sum_{t=2}^{T} (\hat{e}_t - \hat{e}_{t-1})^2}{\sum_{t-1}^{T} \hat{e}_t^2} \tag{21}$$

y_t and x_t are two price series. The null hypothesis of no cointegration is rejected for values of DW significantly different from zero.

2. Dickey Fuller Regression:

$$\Delta\, \hat{e}_t = -\, \phi\, \hat{e}_{t-1} + \hat{\varepsilon}_t \tag{22}$$

where e_t is defined in equation (21) and Δ is the first difference.

Test Statistic : τ_ϕ (the t statistic for ϕ)

This testing procedure considers whether the autoregressive parameter for the estimated residuals from the cointegrating regression (ϕ) is significantly different from one. If there is a unit root of the residuals, then the two series are not cointegrated. The null hypothesis of no cointegration is rejected for values of ϕ which are significantly different from zero. Critical values are provided by Engle and Granger (1987).

3. Restricted VAR:

$$\begin{aligned} \Delta y_t &= \hat{\beta}_1 \, \hat{e}_{t-1} + \varepsilon_{1t} \\ \Delta x_t &= \hat{\beta}_2 \, \hat{e}_{t-1} + \hat{\gamma} \, \delta \, y_t + \varepsilon_{2t} \end{aligned} \tag{23}$$

Test Statistic: $\tau^2_{\beta 1} + \tau^2_{\beta 2}$ (the sum of two t statistics for β_1 and β_2)

This test involves the estimation of a vector error correction mechanism for the cointegrating regression. It bases on the joint significance of the error correction coefficients (β_1 and β_2). This test explains that a cointegrated set of variables can be equivalently expressed as an error correction model in (23). If β_1 and β_2 are jointly significantly different from zero, the null hypothesis of no cointegration is rejected. Critical values are provided by Engle and Granger (1987).

4. Unrestricted VAR:

$$\begin{aligned} \Delta y_t &= \beta_1 \, y_{t-1} + \beta_2 \, x_{t-1} + \hat{c}_1 + \varepsilon_{1t} \\ \Delta x_t &= \hat{\beta}_3 \, y_{t-1} + \hat{\beta}_4 \, x_{t-1} + \gamma \, \Delta \, y_t + \hat{c}_2 + \hat{\varepsilon}_{2t} \end{aligned} \tag{24}$$

Test Statistics: $2[F_1 + F_2]$ where F_1 is the F statistic for testing β_1 and β_2 both equal to zero in (24), and F_2 is the F statistic for testing β_3 and β_4 both equal to zero in (24).

The last test procedure utilizes a vector autoregression which is not constrained on satisfying the cointegration constraints. The null hypothesis of no cointegration is rejected if parameters β_1 and β_2 from (24) and β_3 and β_4 from (24) are jointly significantly different from zero. A failure to reject the null hypothesis indicates the lack of a statistically significant relationship between current changes and past values of the economic variables. It implies a general failure of cointegration between variables (Goodwin and Schroeder, 1991; Engle and

Granger, 1987).

Empirical Results

The results of the cointegration tests are presented in Tables 2 through 4. The results of the pair-wise comparisons of dealer prices and grower prices (Table 2) in different regions are consistent for all four tests and all comparisons. Both dealer prices and grower prices for pinto beans and for great Northern beans are cointegrated across distinct geographic regions. The results of the cointegration tests on dealer versus grower prices are presented in Table 3. The tests are all consistent for pinto beans to reject the null hypothesis of no cointegration between dealer and grower prices. The results for dealer versus grower prices for great Northern beans are somewhat ambiguous. Three of the tests would indicate a rejection of the null hypothesis of no cointegration, but the results of the Durbin-Watson test are that the null hypothesis can not be rejected. The Durbin-Watson test is a weaker test than the other three, so it is more likely that grower and dealer prices for great Northern beans are cointegrated. The results of the tests for cointegration of pinto prices versus great Northern prices in the same region are somewhat ambiguous as well, Table 4. However, three of the four procedures all result in a failure to reject the null hypothesis of no cointegration. It therefore would appear that even though pinto and great Northern beans may be production substitutes, they have distinctly separate markets and are in fact not cointegrated.

In summary, dealer and grower prices are cointegrated across markets for both pinto and great Northern bean prices. Dealer and grower prices also are cointegrated in the same market for both bean varieties. However pinto prices are not cointegrated with great Northern prices for both dealer and grower prices.

Table 2. Cointegration Tests on Dealer and Grower Prices Across Production Regions.

Variety	Market	Test	Dealer			Grower		
			Test Statistic	Critical Value	Decision	Test Statistic	Critical Value	Decision
Pinto	Colorado vs. Idaho	DW	1.669	0.511	Reject Null	1.294	0.511	Reject Null
		DF	10.407	4.070	Reject Null	8.528	4.070	Reject Null
		RVAR	108.056	18.300	Reject Null	110.203	18.300	Reject Null
		UNRVAR	835.308	23.400	Reject Null	599.436	23.400	Reject Null
	Colorado vs. N. Dakota	DW	0.780	0.511	Reject Null	1.658	0.511	Reject Null
		DF	6.278	4.070	Reject Null	10.514	4.070	Reject Null
		RVAR	45.404	18.300	Reject Null	130.033	18.300	Reject Null
		UNRVAR	991.052	23.400	Reject Null	410.364	23.400	Reject Null
	Colorado vs. Nebraska*	DW	2.028	0.511	Reject Null	2.015	0.511	Reject Null
		DF	12.475	4.070	Reject Null	12.383	4.070	Reject Null
		RVAR	171.195	18.300	Reject Null	252.836	18.300	Reject Null
		UNRVAR	19051.684	23.400	Reject Null	1109.706	23.400	Reject Null
	Idaho vs. N. Dakota	DW	0.960	0.511	Reject Null	1.759	0.511	Reject Null
		DF	7.043	4.070	Reject Null	11.059	4.070	Reject Null
		RVAR	51.368	18.300	Reject Null	169.798	18.300	Reject Null
		UNRVAR	433.450	23.400	Reject Null	222.458	23.400	Reject Null
	Idaho vs. Nebraska	DW	1.634	0.511	Reject Null	0.757	0.511	Reject Null
		DF	10.228	4.070	Reject Null	5.989	4.070	Reject Null
		RVAR	95.472	18.300	Reject Null	36.341	18.300	Reject Null
		UNRVAR	27471.768	23.400	Reject Null	2202.996	23.400	Reject Null
	N. Dakota vs. Nebraska	DW	0.778	0.511	Reject Null	1.947	0.511	Reject Null
		DF	6.275	4.070	Reject Null	12.056	4.070	Reject Null
		RVAR	45.334	18.300	Reject Null	247.192	18.300	Reject Null
		UNRVAR	1124.528	23.400	Reject Null	274.156	23.400	Reject Null
Great Northern	Idaho vs. Nebraska	DW	1.058	0.511	Reject Null	0.496	0.511	not Reject Null
		DF	7.417	4.070	Reject Null	4.952	4.070	Reject Null
		RVAR	63.317	18.300	Reject Null	26.610	18.300	Reject Null
		UNRVAR	746.018	23.400	Reject Null	264.242	23.400	Reject Null

Table 3. Cointegration Tests on Dealer Prices vs. Grower Prices for Each Variety in Each Region.

Variety	Market	Test	Test Statistic	Critical Value	Decision
Pinto	Colorado	DW DF RVAR UNRVAR	1.045 7.526 56.418 685.826	0.511 4.070 18.300 23.400	Reject Null Reject Null Reject Null Reject Null
	N. Dakota	DW DF RVAR UNRVAR	2.061 12.810 153.032 307.158	0.511 4.070 18.300 23.400	Reject Null Reject Null Reject Null Reject Null
	Nebraska*	DW DF RVAR UNRVAR	1.017 7.394 75.908 730.282	0.511 4.070 18.300 23.400	Reject Null Reject Null Reject Null Reject Null
	Idaho	DW DF RVAR UNRVAR	1.539 9.820 103.780 452.040	0.511 4.070 18.300 23.400	Reject Null Reject Null Reject Null Reject Null
Great Northern	Idaho	DW DF RVAR UNRVAR	0.517 4.806 38.267 289.882	0.511 4.070 18.300 23.400	Not Reject Null Reject Null Reject Null Reject Null
	Nebraska	DW DF RVAR UNRVAR	0.321 4.086 25.808 295.390	0.511 4.070 18.300 23.400	Not Reject Null Reject Null Reject Null Reject Null

Table 4. Cointegration Tests on Pinto versus Great Northern Prices in the Same Region.

	Market	Test	Test Statistic	Critical Value	Decision
Dealer Price	Idaho	DW DF RVAR UNRVAR	0.089 1.818 4.170 32.044	0.511 4.070 18.300 23.400	Not Reject Null Not Reject Null Not Reject Null Reject Null
	Nebraska	DW DF RVAR UNRVAR	0.072 1.615 3.808 39.64	0.511 4.070 18.300 23.400	Not Reject Null Not Reject Null Not Reject Null Reject Null
Grower Price	Idaho	DW DF RVAR UNRVAR	0.089 1.805 4.327 37.472	0.511 4.070 18.300 23.400	Not Reject Null Not Reject Null Not Reject Null Reject Null
	Nebraska	DW DF RVAR UNRVAR	0.074 1.768 3.585 36.97	0.511 4.070 18.300 23.400	Not Reject Null Not Reject Null Not Reject Null Reject Null

PRICE LEADERSHIP

When prices for each dry bean variety are cointegrated across different regions, it means that prices are moving together closely for each variety in different regions. The next researchable question is: given that prices move together closely is there a region that leads the other regions in establishing price? Theoretically we can apply Granger-causality theorem to verify the lead-lag relationship in dealer prices for each variety in different markets. If any leadership in dealer prices is identified in any selected market, then the grower price in that leading market should also be the price leader since the processors set the grower prices.

Granger Causality tests have been widely applied in previous studies to determine the lead-lag relationship between different series. Examples of previous applications have been the real-trade-weighted agricultural exchange rate and monthly real prices and export sales of crops

(Bradshaw and Orden, 1990), economic growth and defense spending (Joerding, 1986), two wholesale beef price quotes (Faminow, 1981), advertising expenditures and Canadian demand for cheese and butter (Reynolds, et al., 1991), land prices and farm-based returns (Phipps, 1984), and the exchange value of the dollar and the U.S. trade balance (Mahdavi and Sohrabian, 1993).

Theory and Methods of Causality Test

Broadly speaking, a set of variables z_t is said to be cause by x_t in Granger's sense if the information in past and present x_t helps to improve the forecasts of z_t. Suppose one suspects a lead-lag relationship between z_t and x_t, i.e. one suspects that z_t is caused by x_t. One can regress z_t on its own lagged information ($z_{t-1}, z_{t-2}, \cdots$) as well as on past information about x_t ($x_{t-1}, x_{t-2}, \cdots$). If the own lagged information of z_t does not contribute to the prediction significantly (i.e. the past information about x_t is sufficient enough to predict z_t), then one can conclude that z_t is indeed caused by x_t.

Define a regression relationship for the forecast as follows:

$$\begin{aligned} z_t &= \mu_1 + \hat{\alpha}_1 \times z_{t-1} + \hat{\alpha}_2 \times z_{t-2} + \cdots + \hat{\alpha}_q \times z_{t-q} \\ &\quad + \hat{\delta}_1 \times x_{t-1} + \hat{\delta}_2 \times x_{t-2} + \cdots + \hat{\delta}_j \times x_{t-j} \\ &\quad \textit{where } 1 < q < t \quad \textit{and} \quad 1 < j < t \end{aligned} \tag{25}$$

Equation (25) is the "full model" which includes both past x-series ($x_{t-1}, x_{t-2}, \cdots$) and past z-series ($z_{t-1}, z_{t-2}, \cdots$). In order to verify the impacts of $x_{t-1}, x_{t-2}, \cdots, x_{t-j}$ on predicting z_t, a reduced model is defined as:

$$z_t = \mu_2 + \bar{\beta}_1 \times x_{t-1} + \bar{\beta}_2 \times x_{t-2} + \cdots + \bar{\beta}_j \times x_{t-j} \tag{26}$$

Equation (26) is the "reduced model" which includes only the lagged information about x_t. To test the causality from x_t to z_t, apply the "Partial F Test" which can be formulated as the following:

$$\textit{Test Statistic } F = \frac{\dfrac{(\hat{\rho}_{full} - \hat{\rho}_{reduced})}{k}}{\dfrac{(1 - \hat{\rho}_{full})}{1 - p}} \tag{27}$$

where $\hat{\rho}_{full}$ = the R^2 value from the full model

$\hat{\rho}_{reduced}$ = the R^2 value from the reduced model

k = the number of the variables which are included in the full model but not in the reduced model

p = the number of the total variables included in the full mode

The null hypothesis is "lagged z_t information does not contribute significantly in forecasting current z_t". The alternative hypothesis is "lagged z_t information contributed significantly in forecasting current z_t". Accepting null hypothesis means there is a significantly lead-lag relationship between x_t and z_t, i.e., z_t is "Granger caused" by x_t.

Since the edible bean price series are not stationary as shown in the first section, a transformation is necessary to make them stationary series. Taking the first difference of each price series results in stationary series. This has been verified by the Unite Root Tests (SAS/ETS).

Results from the Causality Tests

The results of the Granger Causality tests are displayed in Table 5. Based on the full model and the reduced model, the null hypothesis is "own lagged variable (z_{t-1}) does not contribute significantly when estimating the independent variable (z_t) ". A rejection of the null hypothesis substantiates the alternative hypothesis "own lagged variable (z_{t-1}) contributes significantly when estimating the independent variable (z_t)". If the test statistic indicates to reject the null hypothesis, then there is no significant lead-lag relationship between the two price series z_t and x_t. In this case z_t has to be estimated by both z_{t-i} and x_{t-i}, so there is no significant lead-lag relationship between z_t and x_t. If x_t leads z_t in the Granger's sense, then it is possible to predict z_t using only x_{t-1}, x_{t-2}, ...

The results in Table 5, indicate no lead-lag relationship between the Pinto dealer prices in Nebraska and Colorado because these two series both rely on own lagged prices to predict the future dealer prices. There is also no significant lead-lag relationship between North Dakota and Colorado Pinto dealer prices since they both failed to reject null hypothesis. Idaho appeared to be the only leader in Pinto dealer price based on the Granger test. For Great Northern beans, there is no significant lead-lag results, Table 6. This indicates no lead-lag relationship between the Pinto dealer prices in Nebraska and Colorado because these two series both rely on own lagged prices to predict the future dealer prices. There is also no significant lead-lag relationship between North Dakota and Colorado Pinto dealer prices since they both failed to reject null hypothesis. Idaho appeared to be the only leader in Pinto dealer price based on the Granger test. For Great Northern beans, there is no significant lead-lag relationship between the Idaho and western Nebraska-eastern Wyoming markets. These results do not support our hypothesis that the dominant production region would be the price leader for each bean variety.

DECISION ANALYSIS ON MARKETING MARGINS

It was shown previously that dealer prices are significantly cointegrated with grower prices for Pinto and Great Northern beans in all of the selected markets. This is not surprising since bean processors would adjust bids to growers in response to changing output (dealer)

Table 5. Partial F Tests on Causality for Dealer Prices.

Variety	Assumed Leader	Assumed Follower	Test Statistic
Pinto	Nebraska	Colorado	16.39*
	Idaho	Colorado	2.360
	N. Dakota	Colorado	1.409
	Colorado	Nebraska	13.529*
	Idaho	Nebraska	5.584
	N. Dakota	Nebraska	0.116
	Colorado	Idaho	8.195*
	Nebraska	Idaho	13.842*
	N. Dakota	Idaho	10.401*
	Colorado	N. Dakota	1.081
	Idaho	N. Dakota	3.738
	Nebraska	N. Dakota	0.016
Great Northern	Idaho	Nebraska	0.180
	Nebraska	Idaho	1.077

Note: values with '*' indicate significant at 1% significance level

prices. However, there is no empirical studies documenting how dry bean processors adjust grower prices in response to changing dealer prices. This marketing margin is a subject of concern among bean growers. They perceive that it has widened in recent years beyond any increase in processing costs. In this section, the margin between grower and dealer prices is analyzed and the relative size of the margin is compared to the variability in dealer prices.

<u>Theoretical Framework</u>

Following the framework of Sandmo (1971) and Brorsen, et al. (1985), the theory of marketing margins when decision maker is not certain about the output price is established.

Consider a price-taking firm producing an output y from a raw material input x (in this case, the dry bean crops supplied by local growers) and a vector z of other inputs (capital, labor, etc.). The technology of the firm is represented by the nonstochastic production function

$$y = f(x, z) \tag{28}$$

where f is an increasing and concave function of x and z.

Let p be the output price of y, r be the input price of x, and q be the vector of input prices of z. Lets assume the decision maker knows the input prices r and q at the time of making the decision (Sandmo, 1971), but because of production lags, decision maker does not know with certainty the output price p. Under a random market demand, the probability distribution of output price is determined by the intersection of market supply and demand. Thus considering the randomness of market demand as exogenous to the firm and in the absence of complete contingent markets, p is represented by a random variable with a given probability distribution reflecting the beliefs of the decision maker about output price.

If the decision maker maximizes the expected utility (EU) of the firm's wealth, the firm makes its production decision according to the following model:

$$Max_{x,z}\ E\ U\,[w + p\,y - q'\,z - r\,x \mid y = f(x, z)] \tag{29}$$

Equation (29) defines the firm's wealth as initial wealth (w) plus revenue (p y), then minus cost (q' z + r x). The utility function (U) is increasing and concave for a risk-averse firm, i.e.

$$\begin{aligned} U_w &= \partial U / \partial w > 0 \\ U_{ww} &= \partial^2 U / \partial w^2 < 0 \end{aligned} \tag{30}$$

After solving Equation (29), we will be able to get the risk-responsive input demand (x* and z*) as well as output supply (y*). Each one of them is a function of w, r, q, and the probability function of p. Define p = E(p) and define σ as the second (or possibly higher) moments of the subjective probability distribution of p. We can express y*, x*, and z* as the following,

$$\begin{aligned} y* &= y(w,\ r,\ q,\ \bar{p},\ \sigma) \\ x* &= x(w,\ r,\ q,\ \bar{p},\ \sigma) \\ z* &= z(w,\ r,\ q,\ \bar{p},\ \sigma) \end{aligned} \qquad \textbf{(31)}$$

The properties of the output supply function (y*) have been analyzed in detail by Sandmo (1971) and Ishii (1977). Similarly the properties of the input demand functions (x* and z*) have been discussed by Batra and Ullah (1974), and Hartman (1975).

Following the assumptions made by Brorsen, et al. (1985), we introduce two restrictive assumptions about the production function. First, the production function y = f (x, z) is assumed to be weakly separable, and can be written as y = g [x, h(z)]. This appears to be a reasonable assumption in the dry bean processing industry. Dry bean dealers will take in the raw materials supplied by the growers, then clean them, store them, and export to other regions. Second, we assume the function g has a Leontief or fixed coefficient property such that

$$y = \min\ [\ x,\ h\ (z)\ ] \qquad \textbf{(32)}$$

This means that each unit of the output y requires exactly one unit of x as input. This assumption is vary general and would appear to be most appropriate when the production process involves the transformation or servicing of a commodity (Heien, 1980).

Given the firm's technology is represented by (32), we can solve the optimization problem in (29) by decomposing the maximization into two stages (Batra and Ullah, 1974; Hartman, 1975; Brorsen, et al. 1985):

$$\begin{aligned} & Max_{x,y,z}\ E\ U\ (w + py - q'z - rx) \\ & s.t.\ \ y = \min\ [\ x,\ h(z)\ ] \\ = \ & Max_{y}\ E\ U\ [\ w + py - Min_{x,z}\ (q'z + rx)\ ] \\ & s.t.\ \ y = \min\ [\ x,\ h(z)\ ] \end{aligned} \tag{33}$$

The first stage is a standard cost minimization problem under certainty. Let $x_1(q, r, y)$ and $z_1(q, r, y)$ be the general cost-minimizing input demand functions. It follows from our assumption about the production technology (32) that the indirect cost function has the form for positive prices:

$$C(r,\ q,\ y) = q'z_1 + rx_1 = n(q,\ y) + ry \tag{34}$$

where C is a linear homogenous function, increasing and concave in prices (r, q), and increasing and strictly convex in output y. Furthermore, from the envelope theorem,

$$\frac{\partial C}{\partial r} = x_1 = y \tag{35}$$

$$\frac{\partial C}{\partial q} = \frac{\partial n}{\partial q} = z_1\ (q,\ y) \tag{36}$$

Expression (35) and (36) reflect the implications of the production technology (32) for the specification of the cost minimizing input demand functions x_1 and z_1.

Then the second stage maximization becomes

$$\begin{aligned} &Max_y\ E\ U\ [w + py - rx_1 - q'z_1] \\ = &Max_y\ E\ U\ [w + (p - r)y - q'z_1(q,y)] \end{aligned} \tag{37}$$

If we define M = p - r as the "effective margin", i.e., the difference between the output price and the price of the raw material input, then equation (37) becomes

$$Max_y\ E\ U\ [w + My - q'z_1(q,y)] \tag{38}$$

Equation (38) is an expected utility maximization problem with respect to output y under uncertain margin M. This is similar to Sandmo and Ishii's model except the presence of the margin instead of the output price. The solution of the maximization problem in (38) is the risk-responsive supply function

$$y^* = y(w, q, \bar{p}-r, \sigma) \text{ or } y* = y(w, q, M, \sigma) \tag{39}$$

where p = E(p) and σ is the second (or possibly higher) moments of the subjective probability function of p, as we defined earlier. Brorsen, et al. (1985) indicated that the following relationship exist between the cost-minimizing input demand functions (x_1 and z_1) and the risk-responsive input demand (x* and z*):

$$x^*(w, q, \bar{p}-r, \sigma) = x_1(y^*) = ky^*(w, q, \bar{p}-r, \sigma) \tag{40}$$

$$z^*(w, q, \bar{p}-r, \sigma) = z_1[q, y^*(w, q, \bar{p}-r, \sigma) \tag{41}$$

By total differentiation, it follows that

$$\begin{aligned} \frac{\partial x^*}{\partial(w,\ q,\ \bar{p}-r,\ \sigma)} &= \frac{\partial y^*}{\partial(w,\ q,\ \bar{p}-r,\ \sigma)} \\ \frac{\partial z^*}{\partial q} &= \frac{\partial z_1}{\partial q} + \frac{\partial z_1}{\partial y} \times \frac{\partial y^*}{\partial q} \\ \frac{\partial z^*}{\partial(w,\ \bar{p}-r,\ \sigma)} &= \frac{\partial z_1}{\partial y} \times \frac{\partial y^*}{\partial(w,\ \bar{p}-r,\ \sigma)} \end{aligned} \tag{42}$$

Equation (42) provide some useful information on the properties of the risk-responsive input demand functions. For example, the influence of a change in uncertainty (σ) on the input demand (x^*) will be equal to its influence on output supply (y^*). Furthermore, an increase in uncertainty (σ) will reduce output supply (y^*) means ($\partial y^* / \partial \sigma < 0$), which also implies that a reduction in input demand (z^*) ($\partial z^* / \partial \sigma < 0$) if and only if the inputs in the vector z are noninferior inputs ($\partial z^* / \partial y > 0$).

Following the arguments provided by Brorsen, et al. (1985) and Arrow, it appears to be reasonable to limit our discussion to the case where the firm exhibits decreasing absolute risk aversion (DARA) preferences (i.e., where the Arrow-Pratt absolute risk aversion coefficient $R = -U_{ww} / U_w$ decreases with wealth, and $\partial R / \partial w \leq 0$).

From equation (39) we know that output level y^* is a function of wealth (w), vector output prices (q), effective margin (M), and uncertainty (σ). We can derive the relationship between the effective margin and output level to be the following

$$\hat{M} = \bar{p} - r = M(w,\ q,\ y,\ \sigma) \tag{43}$$

Equation (43) provides the basic theoretical framework for our study. Since we only concentrate in the relationship between margins and price uncertainty, we can simplify expression (43) to be the following:

$$\hat{M} = M(\sigma) \tag{44}$$

Our empirical work is based upon (44), and we will discuss the detail in the next section.

Brorsen, et al. (1985) had a detailed discussion about the relationship between firm's behavior, and its impacts on optimal output level (y*) and other factors (w, q, σ) under price uncertainty. They concluded that given DARA preferences, changes in output level is positively related to the changes in margin. Furthermore, an increase in price uncertainty decreases output, and an increase in price uncertainty always increases the expected margin:

$$\frac{\partial \hat{M}}{\partial y} > 0 \tag{45a}$$

$$\frac{\partial \hat{M}}{\partial \sigma} > 0 \tag{45b}$$

Since we are only interested in the relationship between the margin and the uncertainty, expression (45b) provides the base of the hypothesis in our empirical study.

Data and Procedure

The first thing to be analyzed is to identify the relationship between the margins and price uncertainty for different dry bean varieties in different production regions.

In order to estimate the price risk variable (σ) in equation (44), we need to find a measurement for σ. Previous researchers who have investigated the impact of risk on farmers' production decisions used some measure of past annual price (or income) variability as the measurement of risk, because farmers must make production decisions up to a year before they sell their crop (Just, Lin, Winter and Whitaker). These researchers all used some form of distributed lag or moving average of the deviation of actual price (or income) from expected price (or income). Since dry bean processors may own the dry bean for a shorter period (one or two months, and usually no more than one year), processors should be concerned by the variability of prices over a one-month or two-month period. In our study, we assume that dry bean processors will make a decision on the margin based on the price variability in the previous period. For example, processors will decide the margin for August 1996 based on the price variability in July 1996. We can calculate the standard deviation of each month's dealer price based on our weekly dealer prices, and use this standard deviation as a measurement of the risk which processors will consider. Then we can construct a model to determine the influence of price variability on margins:

$$\hat{M} = \hat{\alpha}_0 + \hat{\alpha}_1 \times T + \hat{\alpha}_2 \times SD_{t-1} \tag{46}$$

where M is estimated margin, T is time period, and SD_{t-1} is the standard deviation of the previous month's dealer price. We incorporate a time variable (T) in (46) because there is an increasing trend in the historical margins for Pinto beans and Great Northern beans in some of the markets (Figure 7 to Figure 12). The increasing trends have also been verified by the Unit Root Tests procedures (Hamilton, 1994).

Once the influence of price variability on the margins is determined, the next step is to forecast the margins for a reasonable period by using the estimated price variability. In order to

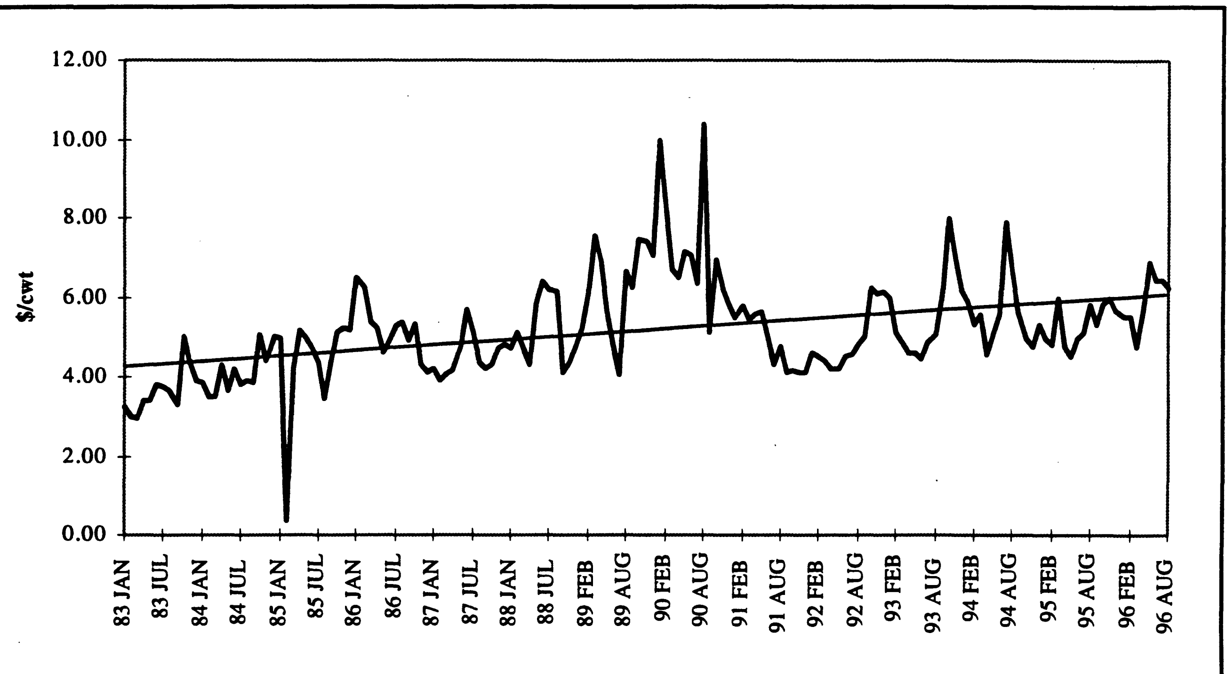

Figure 7. Historical Price Margin for Pinto Beans in Colorado, 1983-1996.

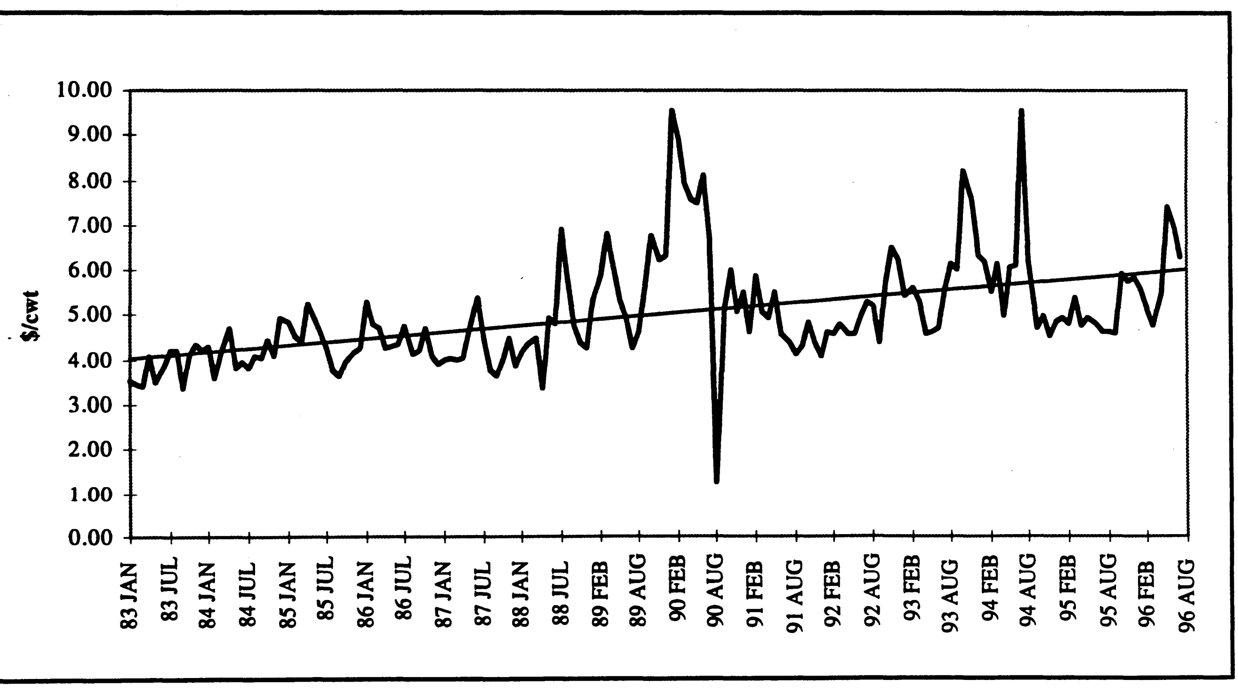

Figure 8. Historical Price Margins for Pinto Beans in Idaho, 1983-1996.

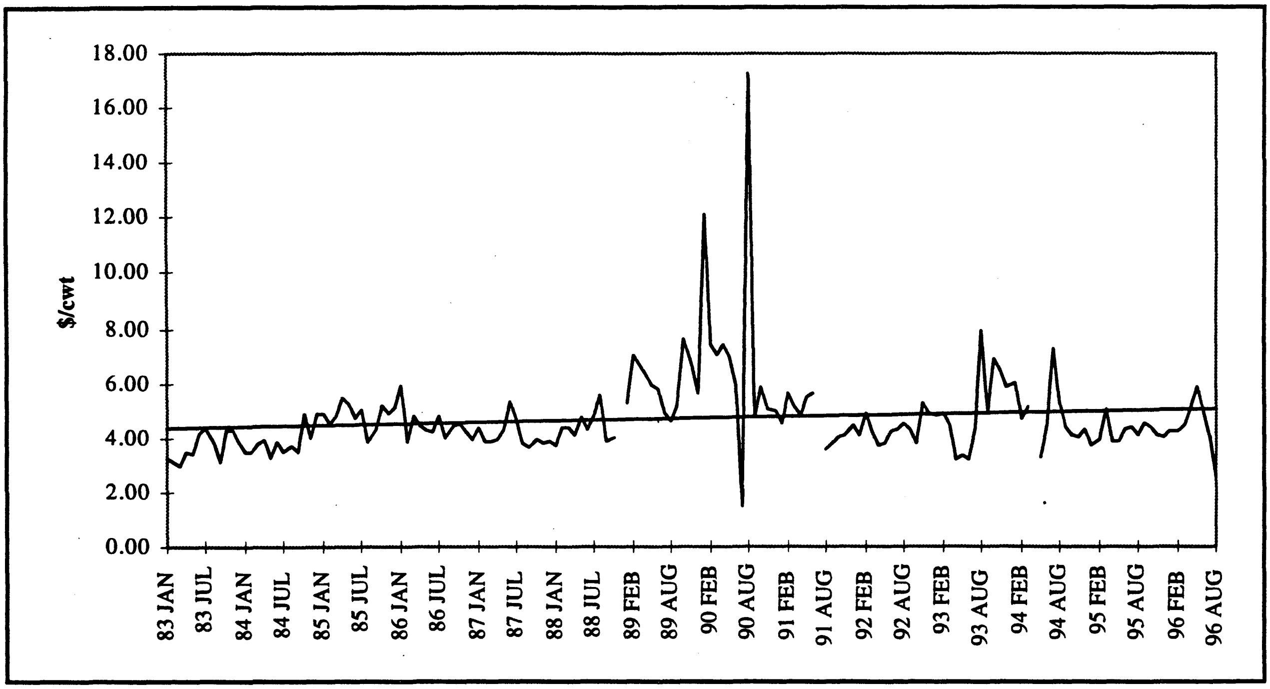

Figure 9. Historical Price Margins for Pinto Beans in North Dakota, 1983-1996.

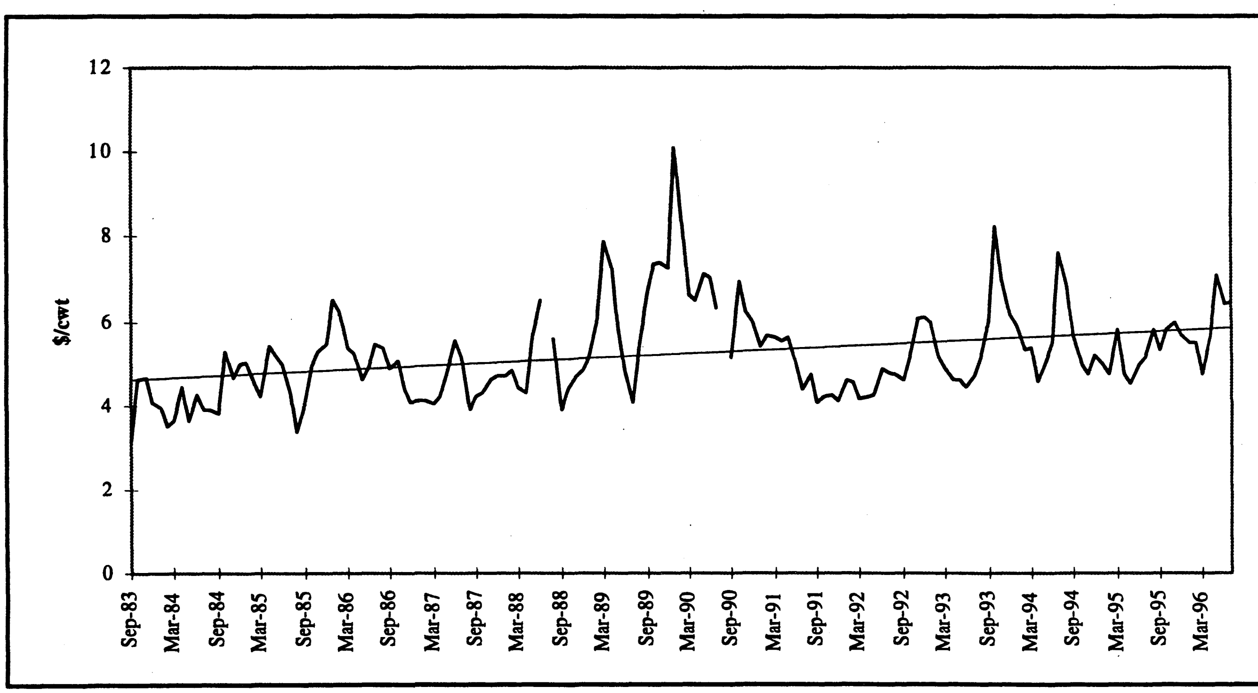

Figure 10. Historical Price Margins for Pinto Beans in Western Nebraska-Eastern Wyoming, 1983.1996.

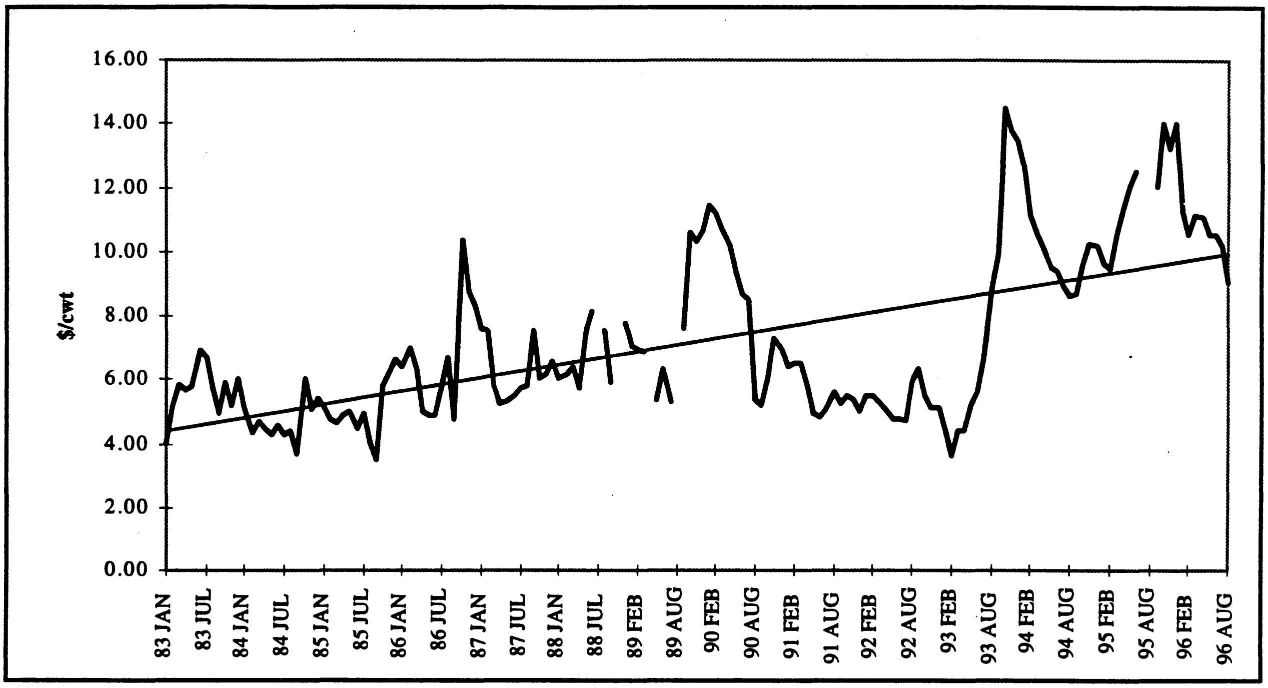

Figure 11. Historical Price Margins for Great Northern Beans in Western Nebraska-Eastern Wyoming, 1983-1996.

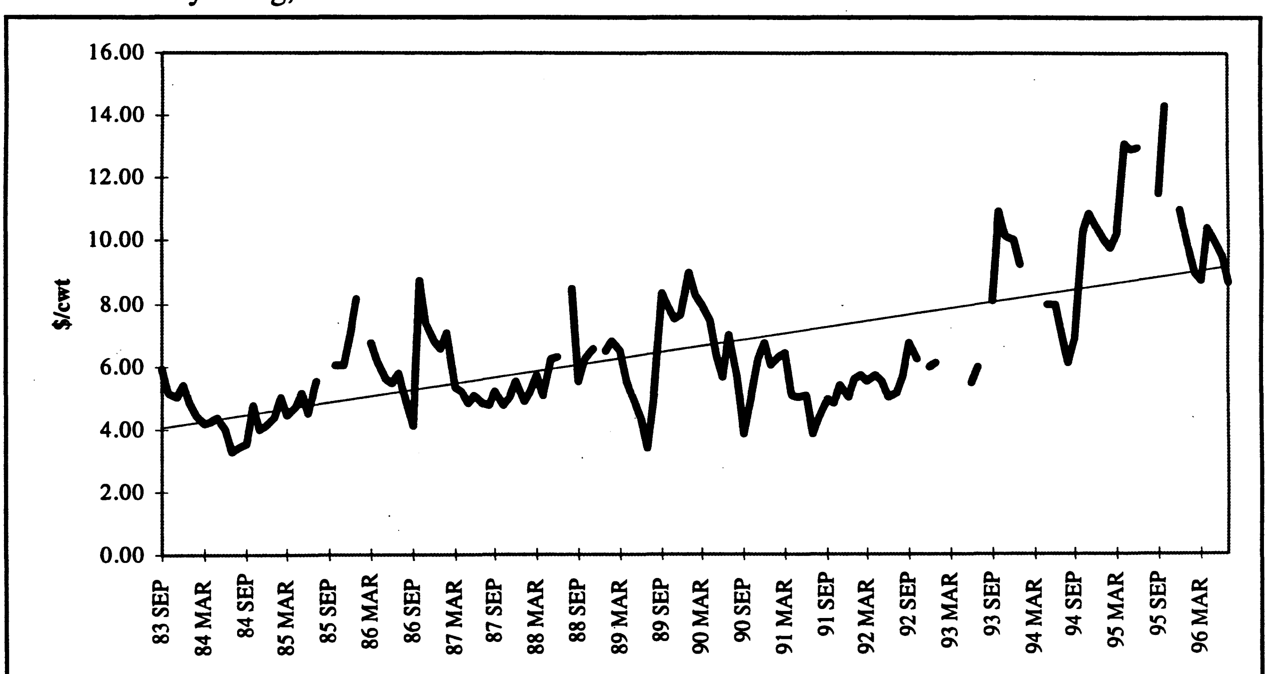

Figure 12. Historical Price Margins for Great Northern Beans in Idaho, 1983-1996.

forecast the margin, an estimate or forecast of the price variability must first be done. Essentially it is a two stage procedure: first, estimate the future price variability; second, forecast the margins utilizing the estimated price variability. There are several methodologies researchers can apply to estimate a time series. An AR(1) process appears to be the most appropriate choice for this data based on the Akaike's Information Criterion (AIC) and Schwarz's Bayesian Criterion (SBC) under the Maximum Likelihood Method (ML).

Assume that the price variability follows the first-order Autoregressive Process or AR(1), then express the standard deviation of dealer price for each month as an AR(1) process which satisfies the following difference equation:

$$SD_t = c + \phi \times SD_{t-1} + v_t \tag{47}$$

where

$$\begin{aligned} E(v_t) &= 0 \\ E(v^2{}_t) &= constant \end{aligned} \tag{48}$$

After estimating the standard deviations of dealer prices for each future month, we can apply the results to forecast the future margins. Based on the Maximum Likelihood Method (ML) and AIC and SBC, it appears that the margin can be best estimated with the following ARMA(1,1) expression:

$$\bar{M}_t = \mu + \omega_0 \times SD_{t-1} + \frac{(1-\theta_1 B)}{(1-\iota_1 B)} \times a_t \qquad \textbf{(49)}$$

where t = indexes time

μ = mean term

B = the backshift operator; i.e. B $X_t = X_{t-1}$

1-θ_1 B = the moving-average operator

1-ι_1 B = the autoregressive operator

More details about the ARMA process can be found in Hamilton (1994) and SAS/ETS User's Guide. Since we are using the ML method to forecast the margins, it is calculated as "unconditional forecasts". Details about unconditional forecasts can be found in SAS/ETS User's Guide.

Results

Table 6 lists the estimated parameters for equation 50. It is obvious that time and previous monthly standard deviation of dealer price have significant influence on the margins. Changes in margins for Pinto beans in Idaho is effected by time most significantly (+0.0086) comparing to Pinto beans in other production regions. Changes in margins are affected by time almost the same for Great Northern beans in both Idaho market (+0.0332), and Western Nebraska and Eastern Wyoming market (+0.0375). But overall changes in margins for Great Northern beans are influenced by time more significantly comparing to Pinto beans.

The results show that previous month's price variability has a positive influence on margins for both Pinto beans and Great Northern beans in all markets (Table 6). This conclusion is consistent with the theoretical hypothesis in (46b): an increase in price uncertainty always increases the expected margin. Among four production regions of Pinto beans, the changes in margin in Western Nebraska and Eastern Wyoming are influenced by the previous month's price

Table 6. Estimated Regression Parameters for Equation: $M_t = \alpha_0 + \alpha_1 \times T + \alpha_2 \times SD_{t-1} + e_t$

Variety	Market	Intercept (α_0)	T (α_1)	SD_{t-1} (α_2)
Pinto	Colorado	4.3525*	0.0086*	0.4105*
	Idaho	4.0851*	0.0093*	0.3504*
	North Dakota	4.5367*	0.0021	0.0863
	W. Nebraska E. Wyoming	4.3363*	0.0069*	0.6192*
Great Northern	Idaho	3.8604*	0.0332*	0.4727*
	W. Nebraska E. Wyoming	3.7938*	0.0375*	1.1772*

Note: The values with '*' indicate significant at the 10% significance level.

variability most significantly (+0.6192), followed by Colorado market (+0.4105), Idaho market (+0.3504), and North Dakota market (+0.0863). For Great Northern beans, the influence of previous month's price variability on the changes in margins in Western Nebraska and Eastern Wyoming (+1.1772) is almost three times as much as in Idaho (+0.4727).

The results of forecasting margins from August 1996 to March 1997 using equation (50) for pinto and great Northern beans are displayed in Tables 7 and 8. The forecasted margins also are compared to the actual margins over the same time period. Model results are quite robust. All but three of the actual monthly margins for Pinto fall with in the 95% confidence interval established with the forecasted monthly margins. All of the actual monthly margins for Great Northern fall within the 95% confidence interval established with the forecasts.

These results would tend to substantiate our initial hypothesis that the margin between grower and dealer prices should be proportional to the variance in dealer prices. The results are also consistent with prior research indicating margins tend to increase with output price uncertainty.

Table 7. Comparison Between Forecasted Margins and Actual Margins for Pinto Beans.

Market	Date	Forecast Margin	Lower 95%	Upper 95%	Actual Margin
Colorado	August 1996	5.92	4.07	7.76	6.13
	September 1996	5.82	3.82	7.82	6.38
	October 1996	5.72	3.62	7.81	7.90*
	November 1996	5.63	3.47	7.79	7.19
	December 1996	5.56	3.35	7.76	7.50
	January 1997	5.50	3.27	7.73	8.25*
	February 1997	5.45	3.20	7.70	7.50
	March 1997	5.41	3.15	7.68	7.13
Idaho	August 1996	5.93	3.46	8.41	6.00
	September 1996	5.78	3.22	8.35	6.81
	October 1996	5.66	3.03	8.29	6.20
	November 1996	5.56	2.88	8.24	7.50
	December 1996	5.47	2.76	8.18	7.17
	January 1997	5.40	2.67	8.13	6.56
	February 1997	5.34	2.59	8.09	6.13
	March 1997	5.29	2.53	8.04	7.50
N. Dakota	August 1996	4.38	1.52	7.24	2.88
	September 1996	4.40	1.51	7.29	4.69
	October 1996	4.42	1.51	7.34	4.70
	November 1996	4.44	1.50	7.38	5.06
	December 1996	4.46	1.50	7.42	4.50

	January 1997	4.48	1.51	7.45	4.88
	February 1997	4.49	1.51	7.48	4.50
	March 1997	4.51	1.51	7.50	3.63
W. Nebraska	August 1996	5.93	4.15	7.71	6.13
E. Wyoming	September 1996	5.56	3.41	7.72	5.63
	October 1996	5.38	3.13	7.62	7.90
	November 1996	5.28	3.01	7.55	7.19
	December 1996	5.23	2.96	7.51	7.50
	January 1997	5.21	2.93	7.48	8.25*
	February 1997	5.20	2.92	7.47	7.50
	March 1997	5.19	2.91	7.47	7.25

Table 8. Comparison Between Forecasted Margins and Actual Margins for Great Northern Beans in Different Markets

Market	Date	Forecast Margin	Lower 95%	Upper 95%	Actual Margin
Idaho	August 1996	8.26	6.21	10.31	6.38
	September 1996	7.94	5.06	10.82	6.88
	October 1996	7.67	4.36	10.99	7.83
	November 1996	4.46	3.87	11.05	7.75
	December 1996	4.28	3.52	11.04	7.83
	January 1997	4.13	3.25	11.00	7.25
	February 1997	7.01	3.06	10.96	6.75
	March 1997	6.91	2.91	10.90	7.00
W. Nebraska	August 1996	10.79	8.74	12.85	8.00
E. Wyoming	September 1996	10.53	7.73	13.34	7.10
	October 1996	10.29	6.98	13.60	9.35

	November 1996	10.06	6.37	13.75	8.31
	December 1996	9.85	5.87	13.83	8.42
	January 1997	9.66	5.44	13.87	8.81
	February 1997	9.48	5.07	13.89	8.69
	March 1997	9.31	4.75	13.88	7.56

CONCLUSIONS

Dry bean prices are recorded in a time-series and the first analysis undertaken of time series is the stationarity of the series. A time series is stationary if its mean, variance, and autocovariances are independent of time. Evident in Figures 1 through 6 of the price series across markets that bean price series are not stationary. Unit root tests in Table 1 corroborate this conclusion. All of the p-values are significant at the .01 level. The null hypothesis of an existence of a unit root for each price series can not be rejected. In other words, none of the dealer price or grower price series are stationary.

The results of the cointegration tests are presented in Tables 2 through 4. The results of the pair-wise comparisons of dealer and grower prices (Table 2) in different regions are consistent for all four tests and all comparisons. Both dealer prices and grower prices for pinto beans and for great Northern beans are cointegrated across distinct geographic regions. The results of the cointegration tests on dealer versus grower prices are presented in Table 3. The tests are all consistent for pinto beans to reject the null hypothesis of no cointegration between dealer and grower prices. The results for dealer versus grower prices for great Northern beans are somewhat ambiguous. Three of the tests would indicate a rejection of the null hypothesis of no cointegration, but the results of the Durbin-Watson test are that the null hypothesis can not be rejected. The Durbin-Watson test is a weaker test than the other three, so it is more likely that grower and dealer prices for great Northern beans are cointegrated. The results of the tests for cointegration of pinto prices versus great Northern prices in the same region are somewhat ambiguous as well, Table 4. However, three of the four procedures all result in a failure to reject the null hypothesis of no cointegration. It therefore would appear that even though pinto and great Northern beans may be production substitutes, they have distinctly separate markets and are in fact not cointegrated.

In summary, dealer and grower prices are cointegrated in across markets for both pinto and great Northern bean prices. Dealer and grower prices also are cointegrated in the same market for both bean varieties. However pinto prices are not cointegrated with great Northern prices for both dealer and grower prices.

When prices for each dry bean variety are cointegrated across different regions, it means that prices are moving together closely for each variety in different regions. Given that prices move together closely is there a region that leads the other regions in establishing price? The results in Table 5, indicated no lead-lag relationship between the Pinto dealer prices in Nebraska and Colorado because these two series both rely on own lagged prices to predict the future dealer prices. There is also no significant lead-lag relationship between North Dakota and Colorado Pinto dealer prices since they both failed to reject null hypothesis. Idaho appeared to be the only leader in Pinto dealer price based on the Granger test. For Great Northern beans, there is no significant lead-lag relationship. These results do not support our hypothesis that the dominant production region would be the price leader for each bean variety.

The marketing margin is a subject of concern among bean growers. They perceive that it has widened in recent years beyond any increase in processing costs. We investigated the margin between grower and dealer prices and the relative size of the margin was compared to the variability in dealer prices.

Following the framework of Sandmo (1971) and Brorsen, et al. (1985), the theory of marketing margins when decision maker is not certain about the output price is established. It is obvious that time and previous monthly standard deviation of dealer price have significant influence on the margins. Changes in margins for Pinto beans in Idaho is effected by time most significantly (+0.0086) comparing to Pinto beans in other production regions (Table 6). Changes in margins are affected by time almost the same for Great Northern beans in both Idaho market (+0.0332) (Table 6), and Western Nebraska and Eastern Wyoming market (+0.0375) (Table 6). But overall changes in margins for Great Northern beans are influenced by time more significantly compared to Pinto beans.

The results show that previous month's price variability has a positive influence on margins for both Pinto beans and Great Northern beans in all markets (Table 7). This conclusion is consistent with the theoretical hypothesis in (46b): an increase in price uncertainty always increases the expected margin. Among four production regions of Pinto beans, the changes in margin in Western Nebraska and Eastern Wyoming are influenced by the previous month's price variability most significantly (+0.6192), followed by Colorado market (+0.4105), Idaho market (+0.3504), and North Dakota market (+0.0863). For Great Northern beans, the influence of previous month's price variability on the changes in margins in Western Nebraska and Eastern Wyoming (+1.1772) is almost three times as much as in Idaho (+0.4727).The results of forecasting margins from August 1996 to March 1997 using equation (50) for pinto and great

Northern beans are displayed in Tables 8 and 9. The forecasted margins also are compared to the actual margins over the same time period. Model results are quite robust. All but three of the actual monthly margins for Pinto fall with in the 95% confidence interval established with the forecasted monthly margins. All of the actual monthly margins for Great Northern fall within the 95% confidence interval established with the forecasts.

These results would tend to substantiate our initial hypothesis that the margin between grower and dealer prices should be proportional to the variance in dealer prices. The results are also consistent with prior research indicating margins tend to increase with output price uncertainty.

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