Role of Delivery Options in Basis Convergence

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The corn futures contract, traded on the Chicago Board of Trade, provides sellers with delivery options about the timing of delivery, the location of delivery, and the grade to be delivered. These options presumably have values that can vary from one delivery month to the next. The joint values of the timing and location options are estimated for each delivery month for the years 1989 through 1997. These estimates are then used in regression models to determine the degree to which they influence basis variability on the first day of the maturity month. Econometric models are also developed to see if the estimated implicit options values are useful in improving the forecasts of basis convergence over the 2-month period prior to maturity. The results suggested that variation in the delivery options values in the corn futures contract does indeed help explain basis variability on the

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first day of maturity. An option-value variable, based on estimated values two months prior to maturity, resulted in occasional, small improvements (from a statistical point of view) in the precision of forecasts. The existence of delivery options increases basis variability at maturity, but it is difficult to use this information to improve forecasts of basis convergence. One limitation of the analysis is that the Chicago cash market had few transactions per day during the sample period, and hence the reported spot prices may be inadequate for making high-quality estimates of the options values. Jrl Fut Mark 22:783–809, 2002

INTRODUCTION

Almost 50 years ago, Working (1953) pointed out that the effectiveness of hedging inventories of grain depended on the predictability of basis convergence. If at the current time t the price of a distant futures contract is above the current cash price and if the two prices converge as the maturity time T approached, then selling futures at t, storing grain, and offsetting the positions at T must earn the amount of convergence. Assuming a frictionless market and a futures contract with no implicit options, the basis for the par commodity will converge to zero at expiration. Thus, a firm hedging of the par asset is exposed to no basis risk and earns the exact convergence; no forecast of convergence is needed.

Delivery options specified in futures contracts and costs of arbitrage result in imperfect convergence; in practice, there is basis risk. However, the basis does not have to converge to zero for hedging effectiveness to be high. What matters is the predictability of the basis change over the storage interval. A standard procedure is to make the change in the basis (over some specified interval) a function of the initial basis in a simple linear-regression model (e.g., Heifner, 1966). The model assumes that the parameters are constant over the sample and forecast periods and that all information relevant for the forecast is contained in the initial basis. Variability in the degree of convergence to date T that is not explained by the size of the initial basis is treated as a random error. The estimated standard deviation of the regression-error term can be used to compute a standard error of forecast that gives a measure of basis risk. For this model, hedging effectiveness depends on a close, stable relationship between the initial basis and the change in the basis over the storage-hedge interval.[1]

[1]Actual delivery can occur until the last business day of the delivery month. The price for deliveries after trading stops is the futures price of the last trading day.

A natural question, therefore, is whether or not it is possible to improve upon the simple linear-regression model as a forecasting tool for basis changes. One possible source of systematic variability in the basis at contract maturity is the varying values of the implicit options in futures contracts. Futures contracts for the grains on the Chicago Board of Trade give the seller alternatives about the timing of delivery within the maturity month, the location of delivery, and the grade of grain to be delivered; and these options presumably have value. Peck and Williams (1991) took the presence of the location and timing options into account in their study of grain markets, by including proxy variables in a regression model. Their proxies, which are possibly helpful in explaining basis behavior, are not observable at the initial date t and hence cannot be used for forecasts.

This article extends the current literature with estimates of delivery option values to help explain basis variability and convergence on the first delivery day. This is justified by the notion that the importance of the delivery options at any time t is measured by their value. The higher the probability that alternate grades/locations may be substituted, the higher the value of the option. Additionally, for any given probability the value of the option increases as the potential gain from exercising the option increases. Thus, both the probability of a change in the relative prices of grades/locations and the magnitude of the change positively affect the value of the option. For any particular grade and location, on the other hand, the two factors increase uncertainty about the basis at maturity—its sign as well as its magnitude. This in turn may negatively affect the predictability of basis changes.

Three models are explored. The first structural model relates to explaining basis variability on the first day of contract maturity. The second structural model represents a component's approach to forecasting the change in the basis over a storage interval. The third model is an empirical approach to predicting basis change by the initial basis and provides a comparative link to previous literature on basis forecasting (Heifner, 1966; Peck & Williams, 1991). The adjusted R^2 of the regressions should increase when the value of the options is included.

Thus, the general objective of this article is to help clarify basis variability at contract maturity. This should improve our understanding of how contract design influences basis variability and may make it possible to better forecast basis variability. The specific application is for the Chicago Board of Trade corn contract (as it existed prior to the changes in delivery provisions in the year 2000).

This article is organized as follows. The first major section reviews the models and data used in estimating option values. Then, the estimated options values for the first day of contract maturity are used as a variable in regression models of basis behavior to see if they improve our understanding of basis variability on this day. Subsequently, the potential for improving forecasts of basis convergence two months prior to maturity is appraised by attempting to estimate options values for this earlier date and using them in models of basis change over the two-month interval. Forecasting both the basis change to the first and the last delivery days is explored.

ESTIMATING OPTIONS VALUES

We first estimate the joint value of the timing option and location options. For corn the timing option gives the seller the alternative of delivering on any business day of the expiration month (the contract, under the specification existing for the sample period, was typically traded during the first 12–16 business days; delivery could be made after trading stopped; our approach treats the last day of trading as the last date of delivery). The location option allows the short to deliver at alternative locations. For corn, this was Toledo, Chicago, and St. Louis; for our sample, the St. Louis option was rarely used and is not included. The quality option allowed delivery of alternate grades of corn at fixed premiums or discounts relative to the par grade of number 2 yellow corn. This option cannot be evaluated because time-series observations are not continuously available on prices of nonpar grades of corn for the various locations.

Estimates of delivery option values are obtained without and then with "convenience yield," and estimates are obtained for each maturity month for the years 1989–1997, inclusive. Although a number of researchers have examined the impact of timing, location, and quality options separately or additively (e.g., Boyle, 1989; Hemler, 1988; Pirrong, Kormendi, & Meguire, 1994; Silk, 1988), the approach used in this article provides a more accurate assessment of the overall effect of delivery options on basis risk. This is because the joint-option estimates improve results by considering the interactive effects of the different options. Although the estimates may still be biased as a result of ignoring the quality option, the bias is likely to be small because the relative prices of deliverable grades of corn usually do not vary much. The following subsections outline the discrete option-pricing model and data used for the estimation. The pricing model is explained more fully in Hranaiova, Jarrow, and Tomek (2001).

Model for Estimating the Value of the Joint Timing and Location Options

In valuing delivery options, the model assumes perfectly competitive markets, no transaction costs, and no taxes. Because we are interested in the value of the options on the first delivery date, the no-transaction cost assumption is perhaps a reasonable abstraction. Most small hedgers offset their positions prior to the expiration month. Thus, traders who potentially might make delivery are likely those with low transaction costs. The other two assumptions are widely used in the theoretical as well as empirical literature on option pricing, but it is true that futures markets become more concentrated as the last day of trading approaches.

The model assumes that storage is costless during the expiration month except for an interest rate that is fixed for each month, but which varies from month to month. Owners of certified warehouses do not have to pay explicit storage charges, and to the extent that they hold futures positions at maturity, other storage costs may be relatively unimportant. Also, with the declining importance of the Chicago spot market, the competition for storage space may be small; it is likely that this is not true for the Toledo location.

In addition, the joint option is analyzed as if no deliveries occur after the last trading day.[2] The futures contract is assumed to allow for the timing and location options only. The location option expands the deliverable supply by extending the set of deliverable locations; in the case of corn, two active locations exist—Chicago and Toledo. The short can deliver at alternative locations and receive a price adjusted according to a given discount/premium schedule. The delivery month runs from $t = 0$ to $t = T$, where T represents the last day of trading in the currently deliverable contract.

[2]For the grain contracts on the CBOT, sellers may tender delivery on the "first notice day" for the expiring contract or on a subsequent business day thereafter. The value of the contract is determined by the settlement price on the date when notice of delivery is given. Thus, a trader could sell a corn contract through the 1:15 p.m. close of trading (noon on the last day of trading) and has until 4:00 p.m. to tender a notice of delivery. The notice of delivery is passed on to a buyer the next day, and delivery and payment occur on day three. Thus, delivery and receipt of payment have a discrete time lag, but the payment is based on the day one settlement price.

The value of the joint timing and location option, $V_{JO,}$ is determined as the difference between the price of a currently deliverable futures contract without the joint option, $F_{w/oJO,}$ and the futures price of the same contract with the option, $F_{w/JO}$

$$V_{JO}(0) = F_{w/oJO}(0) - F_{w/JO}(0)$$

where the values and prices are estimated on the first delivery day and 2 months prior to the delivery month. The no-arbitrage futures price without the delivery options is a martingale under martingale equivalent probabilities, $F_{w/oJO}(t) = \tilde{E}_t[S(T)]$, where $\tilde{E}$ denotes expectation under martingale equivalent probabilities.

To estimate the values of the implicit options, we considered two alternative assumptions about the institutional framework of the delivery process. One assumption is that delivery can take place at the earliest time the next day after entering a futures position, that is, if the contract is entered at time t, the delivery options cannot be exercised until time $t + 1$. The delivery procedure for corn on the CBOT is a 3-day process, but the price received by the seller is determined by the settlement price on the date that delivery is tendered. A second assumption allows delivery on the same day that the position is taken. This is an assumption usually used in the literature. Assumption two is also perhaps consistent with the value of the contract being determined on day one of the delivery process, but as noted subsequently, assumption one provides a convenient way to estimate options values that contain convenience yield. Also, in the statistical models used to analyze basis variability and to forecast basis changes, we found that the options values computed using assumption one gave larger R squares and larger t ratios. In other words, this measure of the options value had better explanatory power, although both measures gave similar qualitative results. On the basis of these practical considerations, the remainder of this article takes as a point of departure that valuing delivery options using the discrete, 1-day lag assumption provides a reasonable approximation.[3]

In an economy with two risky assets (spot commodity in two deliverable locations), the futures price on the last delivery day converges to the lowest of the two spot prices, $\min[S_1(T, s_T), S_2(T, s_T)]$ for all states of the world s_T. On this last delivery day, no option to delay exists, and the short position must be closed by delivery (or offset), yielding a zero cash flow.

[3]The spot price can have seasonal behavior, but this is not important for the intramonth analysis because we are interested in the spot-price process for each contract month separately, that is, the parameters are estimated for each maturity month.

Thus, the only possible nonzero payoff comes from the marking-to-market cash flow. The value of the futures contract at $T - 1$ is 0

$$V_{T-1} = 0 = \tilde{E}_{T-1}\frac{F(T-1; s_{T-1}) - F(T; s_T)}{R}$$

where R is the risk neutral discount factor, yielding the arbitrage free futures price at $T - 1$

$$F(T-1; s_{T-1}) = \tilde{E}_{T-1}[S(T; s_T)]$$

The futures price is determined by backward induction, recognizing that the value of the contract is reset to zero every day through marking to market. Each state of the world in every time period is checked for optimal exercise where the asset is delivered in the cheapest-to-deliver location if delivery is optimal.

The joint option has a character of an American option on a minimum of two assets resulting in the arbitrage-free futures price at any time $0 \leq t < T - 1$

$$\begin{aligned} F(t; s_t) &= \tilde{E}_t[F(t+1; s_{t+1})] - \tilde{E}_t(\max\{F(t+1; s_{t+1}) \\ &\quad - \min[S_1(t+1; s_{t+1}), S_2(t+1; s_{t+1})], 0\}) \end{aligned}$$

At every time period $0 \leq t < T - 1$, the price is lower by the expected value of the joint option in the following period (see Hranaiova, 2000; Hranaiova, Jarrow, & Tomek, 2001 for a more detailed exposition).

The two underlying state variables, that is, the spot prices in the two deliverable locations, are assumed to follow two correlated lognormal processes

$$dS_1(t) = \mu_1 S_1(t)\,dt + \sigma_1 S_1(t)\,dW_1(t)$$

and

$$dS_2(t) = \mu_2 S_2(t)\,dt + \sigma_2 \rho S_2(t)\,dW_1(t) + \sigma_2\sqrt{1-\rho^2}S_2(t)\,dW_2(t)$$

where $dS_k(t)$ denotes a change in the spot price of asset k during a (infinitesimally small) time increment dt, and $dW_k(t)$ is a Wiener process with zero mean and variance dt. μ_k and σ_k represent the drift and volatility of asset k's instantaneous return, and ρ is the correlation coefficient between instantaneous returns on the two assets.[4]

[4]In our empirical analysis, most of the timing-option values for Chicago, with convenience yield, are zero; just 10 of the 45 sample months have positive values. In contrast, 31 of the 45 timing option values are positive for the Chicago data without considering the convenience yield.

A sequence of two-variate, trinomial processes is used to approximate the two-dimensional diffusion process for the spot prices. The model consists of two risky assets and one bond that form a dynamically complete securities market.

The use of martingale equivalent probabilities implies risk-neutral valuation. The risk-free interest rate is assumed constant for the duration of each individual contract month, a reasonable assumption given the short time horizon. This and the fact that the two correlated spot price processes are path independent result in recombining trinomial trees.

Two sets of values are estimated, those without convenience yield and those with convenience yield. Convenience yield is obtained implicitly from the lattice model by assuming that the observed futures prices correctly reflect the value of the joint option. The theoretical futures prices with the options and with no convenience yield are compared with the futures prices observed in the market. By adjusting the value of the convenience yield in the tree-generating process, the theoretical futures price that closely approximates the observed market price is arrived at iteratively. The spot price tree resulting in this futures price is used to estimate the futures price tree without the options. The joint value of the options on the first delivery day is obtained as the difference between the two futures prices at date zero. The "convenience yield" thus captures all the residual benefits and costs of storage.

A positive convenience yield lowers the value of the timing option as it enters the valuation as a negative cost of storage, that is, as a benefit to storage. In theory, convenience yield should be large when current stocks are small relative to the expected new-crop harvest. In this situation, the current futures price is high relative to the price of new-crop futures. A large convenience yield implies a large benefit from holding existing (small) stocks notwithstanding a negative price of storage. The potential return from holding these stocks makes early delivery less desirable.[5]

However, when the location option and the residual estimation of convenience yield are combined, the joint option values increase dramatically. It is possible that the values are inflated through the omission of unobserved delivery costs in the Toledo location. When these costs are not considered, the spot-price data indicate that the payoff from delivering in Toledo is larger than from delivering in Chicago, that is, because the quoted Toledo spot prices are on average lower than in Chicago, even

[5]The Chicago spot market has for the last decade been practically inactive, and the quoted spot prices may not represent transaction prices. Central Illinois average prices may better represent transactions but represent a wrong location. The problem of low-quality spot prices is notorious for agricultural commodities, and is a limitation of this study.

after adjusting for the 3-cent discount for delivery in Toledo, the Toledo basis on the first delivery day suggests that delivery is most often optimal in Toledo. The basis at time t is defined as the futures price minus the spot price at time t. Positive Toledo bases then translate into a negative convenience yield in the model. These large negative convenience yield flows imply high costs of holding the underlying commodity and ascribe a large value to the option to deliver early.

Data

As noted previously, the value of the delivery options is estimated for the corn futures contract traded at the CBOT. Daily data are used for each expiration month (March, May, July, September, and December) from 1989 to 1997. The years before 1989 are influenced heavily by price-support programs and substantial government stocks and are excluded from the analysis. Futures prices are the daily settlement prices. Cash prices are those reported by the USDA for the Chicago terminal market and for Toledo.[6] The prices are reported in ranges, and the midpoint is used as a representative price. The 90-day T-bill rates obtained from the CRSP database are used as risk-free rates. A rate is treated as constant for each delivery month but differs from month to month. The number of trading days in individual delivery months ranges from 12 to 16.

The volatility for each contract month is estimated as a sample variance of the log of daily spot price returns, with the number of sample observations equal to the number of trading days in the individual expiration months (or to that 2 months prior to the delivery months). The parameters are estimated for each month separately and then used to establish the delivery-option value for that particular month. For example, the variance for July 1992 is first estimated and then used to estimate the option value for July 1992. This approach assumes perfect foresight.[7] Martingale equivalent probabilities for the binomial distribution are equal to 0.5 (Jarrow and Turnbull, 1996).

As a result of the concern about the quality of the spot prices (see the footnote 5), two separate sets of estimates are conducted, each using different spot prices. The first series uses the data reported as Chicago cash prices and the second uses Central Illinois cash prices. The Central Illinois prices are on average lower than the Chicago prices, averaging 266 cents per bushel during the years 1989–97 for Central Illinois and

[6]Estimates of implied volatilities can be used instead of the "perfect foresight" volatilities.

[7]Although not reported in this article, the same estimations were conducted using estimates of the timing option only. R^2's increase significantly when the joint value of the timing and location options is used.

274 cents for Chicago, respectively. The initial bases (on the first delivery day) for Central Illinois are often positive and large. The two price series have similar volatilities, although the Central Illinois series is likely more representative of the daily variability in the spot market.

BASIS BEHAVIOR ON FIRST DELIVERY DAY

The estimates of the options values are used in a model to help explain basis behavior on the first delivery day of the contract month. The value of an option captures the importance of the option for that particular month and if found significant in explaining basis variability, would indicate that delivery options could influence hedging effectiveness.

Theoretical models of futures prices assume a single expiration day T and perfect convergence of the futures and spot prices on this date. For any date $t < T$, the futures price in perfect and frictionless markets equals

$$F(t, T) = S(t) \cdot e^{(r-y+c)(T-t)}$$

where $F(t, T)$ is the futures price at time t for a futures contract expiring at time T, and $S(t)$ is the current spot price. r, y, and c denote the interest rate, convenience yield, and storage cost, respectively. This is a result of a cash-and-carry no arbitrage argument, and the model implies $F(t, T) > S(t)$ unless $y > r + c$. The delivery options add a value to the short and ceteris paribus results in a lower futures price

$$F(t, T) = S(t) \cdot e^{(r-y+c)(T-t)} - \mathrm{JO}(t)$$

where $\mathrm{JO}(t)$ is the value at time t of the joint timing and location options. Thus, the basis at time t and location k, $B_k(t)$, is a function of the interest rate, convenience yield, storage cost, time to maturity, and the option

$$B_k(t) = \frac{F(t, T)}{S_k(t)} = e^{(r-y+c)(T-t)} - \frac{\mathrm{JO}(t)}{S_k(t)}$$

The following model for the basis is estimated by OLS

$$\ln B_k = \beta_0 + \sum_{i=1}^{4} \alpha_i D_i + \beta_1 \ln\left(\frac{\mathrm{JO}}{S_k}\right) + \beta_2 \mathrm{FS} + \beta_3 \mathrm{IntRate} + e$$

where $\ln B_k$ is the log of the basis defined as F/S in location $k = 1, 2$, IntRate is the 90-day T-bill rate, FS is the spread between the price of the currently deliverable futures contract and that of the next nearby on day t and is a proxy for storage costs and benefits including convenience

TABLE I

Chicago Basis Behavior Using Joint Option Without Convenience Yield[a]

Variables[b]	*Log Model*		*Linear Model*	
Constant	−0.025 (−1.5)	−0.027 (−1.8)	0.996 (107.7)	0.991 (127.1)
D1 (Mar)	−0.007 (−1.4)		−0.007 (−1.3)	
D2 (May)	−0.007 (−1.4)		−0.007 (−1.4)	
D3 (July)	−0.005 (−0.9)		−0.005 (−0.9)	
D4 (Sept)	−0.007 (−1.2)		−0.006 (−1.1)	
FS	−0.113 (−2.0)	−0.111 (−2.1)	−0.134 (−2.5)	−0.128 (−2.6)
Interest rate	0.056 (0.5)	0.050 (0.5)	0.072 (0.7)	0.062 (0.6)
JO (no CY)	−0.004 (−1.1)	−0.003 (−1.0)	−0.363 (−0.6)	−0.323 (−0.6)
R^2	0.50	0.32	0.36	0.36

[a]Numbers represent regression coefficients and *t* statistics (in brackets).
[b]FS = futures spread, JO = joint timing and location option, and CY = convenience yield.

TABLE II

Toledo Basis Behavior Using Joint Option Without Convenience Yield[a]

Variables[b]	*Log Model*		*Linear Model*	
Constant	0.158 (6.8)	0.159 (7.2)	0.991 (78.8)	0.988 (87.0)
D1 (Mar)	−0.010 (−1.3)		−0.010 (−1.4)	
D2 (May)	−0.008 (−1.1)		−0.008 (−1.2)	
D3 (July)	−0.004 (−0.5)		−0.002 (−0.2)	
D4 (Sept)	−0.013 (−1.6)		−0.013 (−1.7)	−0.008 (−1.5)
FS	−0.153 (−1.8)	−0.170 (−2.2)	−0.117 (−1.6)	−0.126 (−1.8)
Interest rate	−0.253 (−1.6)	−0.248 (−1.6)	−0.278 (−1.9)	−0.306 (−2.2)
JO (no CY)	0.026 (5.3)	0.028 (6.1)	4.421 (6.0)	4.353 (6.1)
R^2	0.64	0.60	0.68	0.65

[a]Numbers represent regression coefficients and *t* statistics (in brackets).
[b]FS = futures spread, JO = joint timing and location option, and CY = convenience yield.

yield, and JO/S denotes the estimated joint option value as a proportion of the spot price. The joint option is in turn defined using the values of the joint option with and without convenience yield, which is distinct from the effect of FS. D_i's are contract month dummies; there are five delivery months, and December is the reference month. The same model is also fitted using a linear functional form, with the observed basis as a function of JO/S_k.

R^2's are relatively high for the regressions reported in Tables I and II.[8] The variables "explain" 50% of the Chicago basis variability and 64% of the Toledo variability (for the full model). When the options values are

[8]Also, combining estimates of the timing and location options is a useful improvement. The use of the joint option improves the fit of the equations.

estimated without convenience yield, this variable has a negative sign with small t ratios for the Chicago location. In contrast, this variable has a positive sign with relatively large t ratios for the Toledo location.

The joint option variable, however, is a significant factor in the "Chicago" regression when Central Illinois spot prices are used to estimate the option values (not shown). Consistent with the results in Table I, the sign of the coefficient is negative when convenience yield is not considered in obtaining the options values. As noted previously, Chicago spot prices may not correctly reflect relevant prices relative to those in Toledo. Central Illinois spot prices are more likely related to actual transactions and are in this sense of better quality. Being in the wrong location (not a direct delivery location), they too cannot yield reliable levels of joint option estimates, but their variability is probably more representative of the actual spot price variability.

Next, the joint option values estimated with convenience yield are substituted for the joint option in the aforementioned regressions. Although the absolute values of the estimates may not be reliable as a result of the low quality of the data, the variability may still be representative of the variability of the actual option values with convenience yield. In addition, the variability of the basis is expected to reflect the variability in the convenience yield as well as costs of delivery.

As the residually estimated "convenience yield" captures the variation in the unobserved costs of delivery as well as the potential benefit from holding the underlying commodity, incorporating this value in the option estimates should enhance their explanatory power in the regression model, and indeed this variable is important statistically in the equations for both delivery locations (Tables III and IV). The positive signs for this variable in both equations, however, are seemingly contrary to the logic of the basis model outlined previously and deserve further discussion. The results reported in Tables III and IV are not a function of a few "aberrant" observations; a clear positive relationship exists between the options values and the basis observations. Thus, the result is a consequence of the price relationships reflected in the data. In both Chicago and Toledo, the empirical reality is that the basis is negative ($F/S < 1$) on the first day of contract maturity in many months (14 of 45 observations for Toledo), whereas the conceptual model implies that this basis will generally be positive (because typically $r + c > y$). In those cases when the spot price is above the futures price, the probability is small that a trader could obtain a benefit from being short futures and long inventory, that is, the value of the embedded options is small. Thus, it follows, not surprisingly, that larger options values are associated with

TABLE III
Chicago Basis Behavior Using Joint Option with Convenience Yield[a]

Variables[b]	*Log Model*		*Linear Model*	
Constant	−0.008 (−1.4)	−0.010 (−2.0)	0.980 (139.3)	0.978 (177.7)
D1 (Mar)	−0.005 (−1.2)		−0.003 (−0.7)	
D2 (May)	−0.007 (−1.5)		−0.003 (−0.7)	
D3 (July)	−0.002 (−0.4)		−0.001 (−0.1)	
D4 (Sept)	0.000 (0.0)		−0.001 (−0.2)	
FS	−0.166 (−4.3)	−0.170 (−5.1)	−0.159 (−4.1)	−0.164 (−5.1)
Interest rate	0.187 (2.0)	0.163 (1.8)	0.170 (1.8)	0.165 (1.8)
JO (with CY)	0.003 (3.3)	0.003 (3.3)	0.025 (3.0)	0.026 (3.5)
R^2	0.50	0.45	0.48	0.46

[a]Numbers represent regression coefficients and *t* statistics (in brackets).
[b]FS = futures spread, JO = joint timing and location option, and CY = convenience yield.

TABLE IV
Toledo Basis Behavior Using Joint Option with Convenience Yield[a]

Variables[b]	*Log Model*		*Linear Model*	
Constant	0.048 (8.0)	0.045 (8.6)	1.005 (110.6)	1.004 (127.2)
D1 (Mar)	−0.007 (−1.4)		−0.001 (−0.1)	
D2 (May)	−0.005 (−1.0)		0.007 (1.1)	
D3 (July)	0.000 (−0.1)		0.004 (0.6)	
D4 (Sept)	−0.008 (−1.5)		−0.010 (−1.6)	
FS	0.097 (2.4)	0.085 (2.4)	0.120 (2.4)	0.113 (2.4)
Interest rate	−0.181 (−1.8)	−0.182 (−1.9)	−0.238 (−1.9)	−0.221 (−1.7)
JO (with CY)	0.010 (11.0)	0.010 (12.0)	0.081 (7.9)	0.081 (7.9)
R^2	0.85	0.83	0.76	0.70

[a]Numbers represent regression coefficients and *t* statistics (in brackets).
[b]FS = futures spread, JO = joint timing and location option, and CY = convenience yield.

positive bases. Although our findings seem inconsistent with the simple theoretical model, the estimated values of options in futures contracts do help explain basis variability.[9] Like most research, our results are a step along the path of increasing knowledge.

Although not the central focus of this research, the use of the joint option values tends to increase the *t* ratios of the futures spread and interest-rate variables, but the signs and magnitudes of the coefficients

[9]The assumptions made about the spot-price generating process, although acceptable for a 12–16 day period, are more problematic over a 2-month period. These prices likely have seasonality and mean-reverting behavior, and a natural extension of our research would be to consider an alternative model of spot prices. Recall, however, the parameters for the price processes are estimated separately for each maturity month.

tend to be unstable across alternative models. The seasonal dummy variables are not important statistically, and this is likely a consequence of the seasonality in the estimated option values.

BASIS FORECASTS

Another way of looking at basis risk involves predictability of basis changes. A good forecast of convergence of the spot and futures prices would be helpful to commercial users of futures. For example, consider storage and hedging decisions, using a May futures contract for corn whose price converges to within 5 cents per bushel of the local spot price at the expiration date. Determining how good this convergence is depends on the basis at the time when the initial futures position was established. If, for example, the basis for the May contract was 15 cents on March 1, the convergence of 10 cents likely provides a satisfactory return, but if the contract's basis was 8 cents on March 1, the convergence of 3 cents will not cover the costs of storage. Thus, it is the change in the basis relative to the initial basis that is relevant for hedgers' storage decisions. Moreover, this change must be predictable. If the 15-cent basis on March 1 had converged to 10 cents, rather than 5, the hedge would not have earned a positive return.

We consider two basis-forecasting models. First, in a structural model, daily changes in the basis are decomposed into separate components and then aggregated by summing over time from the initial date 0 to maturity T. This approach permits the analysis of the impact of individual basis components on the predictability of basis change.

The second model makes the change in the basis a function of the initial basis. This approach has been used in previous studies of grain futures contracts (e.g., Heifner, 1966; Peck & Williams, 1991; Working, 1953). Peck and Williams (1991) have discovered that during the crop years 1964–65 through 1988–89, the initial basis predicted the basis change to the end of trading better than to the first delivery day date in the maturity month, that is, their results indicate that basis risk on the last delivery day of an expiration month is lower than that on the first delivery day. Their model uses proxies for delivery options that measure the actual distribution of the timing and location of deliveries for each particular month. The values of these variables are not known on the initial date when forecasts are made (two months before the expiration month), and hence cannot be used in a forecasting model.

As an alternative, we use estimates of delivery option values at the initial date 0 in models examining the impact of the options on basis risk as well as on the predictability of the change in the basis. The models are evaluated for convergence to the first delivery day and to the last trading day in the expiring contract. The joint value of delivery options two months before the expiration month is estimated for each of the 45 expiration months included in the sample. The methods are described subsequently.

The initial date for basis forecasting was chosen on the basis of observed practices of market participants and for computational convenience. Thus, hedgers are assumed to hold positions for two months that may be rolled forward as a contract expires, and hence offset their position in the expiring contract and establish a new position in the next nearby contract on the first delivery day. They are, therefore, concerned about basis risk at the rollover time, which is assumed for convenience to be the first day of the expiration month. Results are also reported for basis change to the last trading day as in Peck and Williams (1991).

One of the modeling issues, in the applications that follow, is whether or not to pool the observations for the five maturity months. Typical forecasting models for basis changes have fitted separate, simple linear-regression models for individual delivery months (as in Heifner, 1966). However, we have only nine observations per month and consider models with multiple regressors. Thus, alternate levels of aggregation across months are considered. One approach pools all 45 observations for the five contract months. A second approach splits the sample into a December-March-May group (27 observations) and a July-September group (18 observations). This is justified by considering July and September as months in which prices are affected by the transition from the old to the new crop year. Finally, we explored fitting equations for the separate delivery months (nine observations per equation). Although we characterize the range of results in the text, the tables present only a subset of the results.

Structural Model

To see the impact of the individual components of the basis on basis change predictability, the basis is decomposed as follows

$$
\begin{aligned}
B(t) &= F(t, T) - S(t) \\
&= [S(t) + \mathrm{CS}(t, T) - \mathrm{CY}(t, T)]e^{r(T-t)} - \mathrm{JO}(t, T)e^{r(T-t)} - S(t)
\end{aligned}
$$

where $S(t)$ is the spot price at initial time t, $F(t, T)$ is the futures price at time t of a futures contract with expiration at T, CS(t, T) is the present value of the cost of storing for the period of time $(T - t)$, CY(t, T) is the present value of the convenience yield, JO(t, T) is the joint value at time t of the delivery options, and r is riskless interest rate. The notation is simplified according to the following: MCS(t) = MCS = marginal cost of storage, MCY(t) = MCY = marginal convenience yield, both considered constant, and JO(t, T) = JO(t). Taking first differences yields

$$\Delta B(t) = B(t+1) - B(t) = [S(t+1)e^{-r} - S(t)]e^{r(T-t)} + (\mathrm{MCS} - \mathrm{MCY}) - [\mathrm{JO}(t+1)e^{-r} - \mathrm{JO}(t)]e^{r(T-t)} - \Delta S$$

where Δ denotes daily changes. Summing over time from inception to maturity, an aggregate change in the basis is obtained

$$\begin{aligned} B(t) - B(0) &= \sum_{t=0}^{T-1} \Delta B(t) \\ &= S(0)(1 - e^{rT}) - (\mathrm{MCS} - \mathrm{MCY})T \\ &\quad - [\mathrm{JO}(T) - \mathrm{JO}(0)e^{rT}] \end{aligned}$$

Testing the preceding linear model of the aggregate change in the basis amounts to testing the joint hypothesis of the model being correct and that of hedging effectiveness. The following linear model is estimated:

$$B(T) - B(0) = \beta_0 + \beta_1 S(0)(1 - e^{rT}) + \beta_2 \mathrm{FS} \cdot T + \beta_3[\mathrm{JO}(T) - \mathrm{JO}(0)e^{rT}] + \varepsilon(T, 0)$$

where T denotes the number of days between the initial day and the terminal day if entering as a multiplier and the terminal date if entering as a subscript. The terminal date is in turn defined as the first delivery day and the last delivery day. FS denotes the futures price spread, that is, the difference between the price of the current contract, and the next nearby at date t and is used as a proxy for storage and opportunity costs as well as benefits of holding the commodity (including convenience yield).

If the model is correct, we should obtain $\beta_0 = 0$, $\beta_1 = 1$ and $\beta_2 = \beta_3 = -1$. Coefficients different from those expected would indicate problems with the model specification or with measurement errors (e.g., errors in estimating convenience yield and imprecise transaction data).

The preceding regression is location specific, with the dependent variable being the basis change from the first of the month two months before the delivery month to the first delivery day and then to the last delivery day. For example, the Chicago basis change for the May 1992 contract is measured as the difference between the initial Chicago basis on March 1, 1992 and the Chicago basis on May 1, 1992. $S(0)$ is the location specific spot price on the initial day (the same spot price used to calculate the initial basis), and r is the interest rate on the initial day, assumed constant for the duration of the contract. $JO(T)$ is the value of the joint option on the first delivery day, as previously estimated, and $JO(0)$ is the value of the joint option at the initial time 0. As the initial date is two months before the expiration month, the pricing model for the joint option value on the initial day is a combined European and American option, that is, the joint option is a European option for the first two months until the first day of the expiration month. This is the first day when the option can be exercised. From then until the last trading day in the expiring contract, the joint option has a character of an American put.

The inputs for the option pricing model, $S(0)$ and r, are as measured on the initial day 0. The volatility parameter is calculated as the standard deviation of the log of daily returns for the whole period and hence is also assumed constant. Computational methods are analogous to the ones previously described and elaborated on in Hranaiova (2000).[10]

The results from the structural model of aggregate change of basis in Chicago and Toledo are reported in Table V. The data are pooled for all maturity months in the sample. Two horizons are considered for the basis change—one from the initial date 0 (two months before the delivery month) to the first delivery day, and the other one spans from the initial date 0 to the last trading day.

The joint option variable is, with one exception, insignificant in the models of the aggregate changes in the Chicago and Toledo bases. The exception uses the joint option without a convenience yield variable in the Toledo basis change to the first delivery day. Also, the fact that none of the coefficients β_1, β_2, and β_3 equal one indicates measurement errors and/or model-specification problems. If the data are disaggregated into the December-March-May and July-September groups, the coefficients change, but the same qualitative conclusions are reached.

[10]A t value of 1.2 in a model with 24 degrees of freedom is associated with a P value of approximately 0.25. We use a t ratio of one as a benchmark because ratios larger than one in absolute value guarantee that the standard error of the error term will decline and the R^2 corrected for degrees of freedom will increase. We do not believe that arbitrary levels of statistical significance, such as the 0.05 level, are appropriate for the purposes of this research, but readers can obtain the relevant significance levels, if they so choose, by using the reported t ratios.

TABLE V

Structural Model for Predicting Change in Basis with Delivery Options Estimated Without Convenience Yield

	Change of Basis to the 1st Delivery Day[a]		*Change of Basis to the Last Delivery Day*[b]	
Chicago	*Without CY*[c]	*With CY*	*Without CY*	*With CY*
Constant	8.305 (1.2)[e]	7.630 (0.9)	−6.628 (−0.9)	−3.735 (−0.4)
Spot price	3.403 (1.3)	3.376 (1.2)	−3.857 (−1.4)	−3.381 (−1.1)
FS[c]	−0.569 (−3.9)	−0.606 (−4.2)	0.934 (6.0)	1.001 (6.4)
Joint option[d]	1.761 (0.7)	0.020 (0.4)	−2.618 (−1.0)	−0.050 (−1.0)
R^2	0.46	0.46	0.64	0.64
	Change of Basis to the 1st Delivery Day[a]		*Change of Basis to the Last Delivery Day*[b]	
Toledo	*Without CY*[c]	*With CY*	*Without CY*	*With CY*
Constant	12.229 (2.0)	10.315 (1.3)	−14.861 (−1.6)	−9.181 (−0.8)
Spot price	4.439 (1.8)	3.728 (1.4)	−6.371 (−1.8)	−4.935 (−1.3)
FS[c]	−0.548 (−4.0)	−0.482 (−3.4)	0.729 (3.6)	0.734 (3.6)
Joint option[d]	−4.417 (−1.9)	−0.011 (−0.2)	1.958 (0.6)	−0.045 (−0.7)
R^2	0.44	0.39	0.41	0.42

[a]Change from the initial date (1st of the month two months prior to the expiration month) to the first business day of the expiration month.
[b]Change from the initial date to the last trading day in the currently deliverable contract.
[c]FS = futures spread as a proxy for costs and benefits of holding the commodity.
[d]Option value estimated at the initial date.
[e]*t* Ratio in parentheses.

Initial Basis Model

An alternative model makes the basis change a function of the initial basis level and a variable that measures the option value, both of which are observed two months prior to the expiration month. For example, for a May futures contract, the initial basis on March 1 together with the value of delivery options on March 1 are used, first, in models of the change in the basis from March 1 to the first delivery day (say, May 1), and second, in models of the change in the basis from March 1 to the last day of trading (say, May 20). The standard of comparison is the simple linear regression that omits the option value variable. From a statistical viewpoint, does the option value variable add significantly to the explanatory value of the model?

As noted previously, various levels of pooling of observations are considered in estimating the models. The alternatives considered are the following: (1) pooling all 45 observations in a single equation, (2) splitting the sample into two groups, December-March-May and July-September, and (3) five individual months. For the pooled data, equations were fitted with and without monthly dummy variables. Thus, the full regression model for the pooled observations is

$$B(T) - B(0) = \beta_0 + \sum_{i=1}^{4} \alpha_i D_i + \beta_1 B(0) + \beta_2 \mathrm{JO}(0) + \varepsilon(T, 0)$$

where $B(T)$ is defined as the basis of the currently expiring contract on the first (or last) day of trading, and $B(0)$ is the basis two months earlier. Results are obtained for three locations, and in each case, two alternative measures of option values are considered, the joint timing-and-location option with and without an imputed convenience yield.

The results reported in Tables VI and VII are for the models fitted to the pooled December-March-May data set. Results for the pooled 45

TABLE VI

Regressions for Basis Change to First Day, December–May Contracts

	Intercept	*Initial Basis*	*Option Value*	*Adjusted R^2*	*Regular Statistical Error*[b]
Chicago					
No option	−2.702	−1.051	—	0.644	3.583
	(3.35)[a]	(6.93)			
With JO	1.540	−1.025	−2.024	0.666	3.470
	(0.57)	(6.94)	(1.63)		
With JOwCY	−0.851	−0.931	−0.433	0.650	3.553
	(0.49)	(5.14)	(1.19)		
Central Illinois					
No option	10.878	−1.398	—	0.744	3.991
	(5.40)	(8.76)			
With JO	12.908	−1.388	−1.754	0.748	3.962
	(4.87)	(8.75)	(1.17)		
With JOwCY	13.291	−1.066	−1.156	0.766	3.817
	(5.68)	(4.49)	(1.82)		
Toledo					
No option	2.132	−0.789	—	0.791	3.107
	(2.38)	(9.98)			
With JO	6.456	−0.637	−2.630	0.799	3.046
	(2.04)	(4.81)	(1.42)		
With JOwCY	−3.010	−1.355	1.966	0.819	2.894
	(1.21)	(5.05)	(2.19)		

[a]Number in brackets represent *t* statistics.

[b]In cents.

TABLE VII
Regression for Basis Change to Last Day, December–May Contracts

	Intercept	*Initial Basis*	*Option Value*	*Adjusted R^2*	*Regular Error*[b]
Chicago					
No option	−1.794	−1.243	—	0.532	5.313
	(1.50)[a]	(5.53)			
With JO	0.952	−1.226	−1.310	0.522	5.371
	(0.23)	(5.36)	(0.68)		
With JOwCY	0.506	−1.094	−0.538	0.532	5.315
	(0.19)	(4.04)	(0.99)		
Central Illinois					
No option	8.417	−1.224	—	0.581	5.023
	(3.32)	(6.09)			
With JO	12.268	−1.205	−3.328	0.617	4.802
	(3.82)	(6.26)	(1.83)		
With JOwCY	11.495	−0.801	−1.475	0.618	4.795
	(3.91)	(2.68)	(1.85)		
Toledo					
No option	2.458	−0.953	—	0.774	3.948
	(2.16)	(9.49)			
With JO	0.454	−1.024	1.219	0.767	4.008
	(0.11)	(5.88)	(0.50)		
With JOwCY	1.180	−1.094	0.489	0.766	4.016
	(0.34)	(2.94)	(0.39)		

[a]Numbers in brackets represent *t* statistics.
[b]In cents.

observations are available in Hranaiova and Tomek (2001). Table VI pertains to the basis change to the first day of the maturity month, and Table VII portrays the basis change to the last trading day of the month. In both tables, the results are for the equations that exclude the monthly dummies. The dummy variables are not statistically important, and their exclusion has little effect on the magnitudes of the coefficients of the remaining variables.

In interpreting the results, the reader is reminded that the basis is in cents per bushel. For Chicago, the slopes of the initial basis variable are approximately −1.0 (Table VI), implying a one-to-one relationship between the size of the initial basis and its change to the first day of maturity. Also, it is logical that the intercept coefficients of Central Illinois regressions are positive and statistically important because these equations pertain to a nonpar delivery point.

The options value coefficients reported in Table VI all have *t* ratios larger than one. Thus, the R^2 corrected for degrees of freedom increases, and the standard error of the error term declines when either definition of this variable is added to the model. The option value variable has a consistent but statistically modest effect in explaining basis convergence for the three locations considered. The same qualitative result exists when all of the observations are pooled, but typically not for the July-September data set.

The results are mixed for the individual months' regressions. In the majority of months, the options variable is not statistically important, but the variable with imputed convenience yield has t ratios larger than two for March and May in Central Illinois and for March and July in Toledo. The options variable also has t ratios larger than one in seven other equations. In sum, in 11 of 30 comparisons, the options variable's t ratio is larger than one. Because the individual month regressions have only six degrees of freedom, it is possible that the sample data often have insufficient information to ascertain the possible significance of the options variable, that is, it is unclear whether the option variables truly have a zero parameter (cannot reject the null of zero effect) or whether the information in the short sample is inadequate to ascertain the possible significance of the variable. However, as noted, the variable does appear to be important in some cases.

For basis changes to the last day of trading, the options value variable is not statistically important in the Chicago and Toledo equations (Table VII). In contrast, the t ratios are over 1.8 for the Central Illinois equations. Perhaps there are "last day" effects in the spot prices at the par delivery locations that do not exist at the nonpar location, but we do not have a good reason for the contrasting results. In any case, the general point is that our results do not confirm the results of Peck and Williams (1991) (that the initial basis has a larger explanatory power for forecasting the basis change to the last day of trading). If the results in Tables VI and VII are compared, the equations for the basis change to the first delivery date have larger R^2's than do those for the last day for the comparable models.

From the viewpoint of most hedgers, interest would center on forecasting the change in basis to the first day of contract maturity. If the hedge were to be continued past this date, the usual practice would be to roll over the futures position prior to the first day of delivery. Thus, we appraise the potential forecasting value of equations fitted to basis changes to the first day of contract maturity. To do this, we use the median and maximum initial basis observed within the sample for each location (and the corresponding estimated options values) to compute point forecasts of basis change. These are, of course, "ex post forecasts" using the options values that can be estimated from existing data. To evaluate

TABLE VIII
Basis-Change Forecasts to First Day, December–May (Cents)

Model	*Forecast of Maximum*	*t Ratio*[a]	*Forecast of Median*	*t Ratio*
Chicago				
No option	−16.628	4.177	−6.381	1.748
With JO	−16.708	4.333	−5.037	1.387
With JOwCY	−17.067	4.305	−5.425	1.463
Central Illinois				
No option	−19.528	4.468	−6.596	1.622
With JO	−19.016	4.361	−5.658	1.375
With JOwCY	−19.636	4.696	−4.900	1.225
Toledo				
No option	−14.435	4.352	−4.771	1.508
With JO	−12.978	3.806	−4.330	1.389
With JOwCY	−13.850	4.466	−5.819	1.949

[a]Ratio of point forecast to standard error of forecast.

the effect of including the option variable in the model, we report the ratio of the point forecast to its standard error of forecast. Comparisons of these t ratios provide a measure of the precision of forecasts that does not depend on the magnitude of the individual forecast. Recall, the standard error of forecast varies with the magnitudes of the regressors; the computational method used is from Greene (2000, p. 207).

Results based on the pooled December-March-May data are reported in Table VIII. These are the models in which the options value variables were added modestly but consistently to our ability to explain the variability of the basis change to the first day of delivery. From a forecasting perspective, the results first suggest that the addition of the option value variables to the model typically has a small effect on the magnitudes of the point forecasts. If a decision maker were comparing the point estimates (only), then the alternate models would not likely influence the decision.

Second, the joint-option-with-convenience-yield variable increased the precision of forecasts made for the maximum basis level in all three locations, but only in Toledo for the forecast made at the median basis level. With one exception, the joint option (without convenience yield) did not improve the precision of forecasts for those made at either the maximum- or median-level basis levels. Adding an options variable to the model may improve the precision of forecasts, but clearly the improvement, when it exists, is small for the pooled regression models.

TABLE IX
Regression for Basics Change to First Day, May Contract, Central Illinois

Variable	*Model 1*	*Model 2*
Intercept	21.107	27.532
	(5.37)[a]	(8.11)
Initial basis	−2.163	−1.539
	(6.56)	(5.09)
JOwCY	—	−2.668
		(3.05)
$\bar{R}^2$	0.787	0.902
Regression statistical error	3.951	2.672
Median forecast	−5.925	−5.843
	(1.42)	(2.06)
Maximum forecast	−9.169	−9.500
	(2.15)	(3.29)

[a]Ratios shown in parentheses.

As noted previously, the addition of an options variable to the models for individual months did significantly improve the explanation of basis convergence to the first of the month in four cases. The implication for forecasting is demonstrated for one of these cases in Table IX, namely, for the basis change from March 1 to May 1 for the Central Illinois location. In this case, the option variable (with imputed convenience yield) has a *t* ratio of 3.0. Like the results in Table VIII, the point forecasts for the two models differ little, but in this case, the addition of the options variable improves the precision of the forecast considerably. Analogous results were obtained for the March contract for the Central Illinois basis and for the March and July contracts for the Toledo basis.

The results for individual months and locations suggest that the inclusion of an options value variable has the potential to improve the precision of forecasts of basis convergence. However whether or not this is indeed true more generally must await additional empirical analysis that involves more observations and contracts for other commodities. Also, a problem in any realistic forecast is the difficulty of obtaining the ancillary estimates of the joint option values two months prior to maturity.

As indicated in Table VII, the addition of option value variables did not help explain the variability in basis convergence to the last day of trading for the Chicago and Toledo locations. In contrast, both estimates of options values had *t* ratios over 1.8 for the Central Illinois location. This improved explanatory power did not, however, translate into large improvements in the precision of forecasts. Using the equations based on the pooled December-March-May estimates, the t ratio for the point forecast on the basis of the maximum initial basis was 4.47 for the simple linear regression and 4.70 for the comparable forecast using the JOwCY variable. There was no improvement in precision for forecasts made at the median initial level of the basis.

In summary, the initial basis two months prior to maturity provides a better explanation of convergence to the first day of the maturity month than to the last day of trading in the maturity month. Adding option value variables often gives small improvements in the explanatory power of models of basis convergence to the first of the month, but this improved explanatory power does not always translate into more precise forecasts of convergence. Thus, the evidence suggests that the existence of options in futures contracts is a partial explanation of basis variability, but forecasts from such equations do not necessarily reduce basis risk. The options variable was most helpful when the options value was based on an implicit convenience yield and when the forecast was conditioned on a large initial basis.

CONCLUSIONS

Values of delivery options were estimated for the Chicago Board of Trade futures contract for corn for the years 1989–1997. These estimates were then used in statistical models of basis behavior at contract maturity. The variable joint-option-with-convenience-yield had a coefficient with a large t ratio in both the Chicago and Toledo basis equations. When convenience yield was not considered, the option value variable remained statistically important in the Toledo equation but not in the Chicago equation. If, however, Central Illinois price data were used to estimate the option values, then the variable remained important in the Chicago equation.

Therefore, variability in the basis on the first delivery day is partly explained by the varying values of the implicit options in the contracts. However, the coefficient estimates are not very stable. Also, results are difficult to interpret. Specifically, based on the reported price data, the largest share of deliveries should have occurred in Toledo; it typically appears to have been the least cost-delivery location. In fact, the largest proportion of deliveries occurred in Chicago. Presumably, delivery costs (or benefits) exist in these two markets that are not captured by the prices available to us. Because the estimated options values reflect the reported prices, which may not fully reflect the costs of delivery in the two locations, our estimated options values and subsequent regression estimates may be biased.

It is nonetheless of interest to see if the estimated options values can improve forecasts of basis convergence, and two models are estimated for this purpose. The first is a structural model approach to forecasting. This model, when fitted to the full sample, typically had moderate R^2's and small t ratios for the options value variable. Fitting the model to subgroups of the data usually resulted in even smaller R^2 coefficients. This approach to forecasting was clearly unproductive.

The second model builds on Working's (1953) insight that the change in the basis from an initial date until contract maturity can be forecasted from the initial basis by adding an options value variable to his model. Thus, having estimated implicit option values for the corn futures contract at two months prior to maturity, this variable is added to simple linear-regression equations. This variable often improves the statistical fit of the model, especially for the case when the change in basis is defined to the first delivery day. If the estimates of the options values allow for convenience yield, the improvement in statistical fit is usually better.

Hedging effectiveness, however, depends on basis risk, and for merchants storing corn, this depends on the ability to forecast basis convergence over the life of the hedge. Thus, the question is whether the inclusion of an options value variable improves forecasts. Our results suggest that the inclusion of such variables has little effect on the magnitude of the point forecasts and that the precision of forecasts is improved by a modest amount in some cases. However, the small improvement in explanatory power, when it occurs, does not always translate into improved precision in forecasting (i.e., smaller standard errors of forecast relative to the magnitude of the forecast).

Some exceptions are notable in our analysis. For selected models of individual months, the precision of forecasts was improved considerably by the inclusion of an options value variable. In general, the improvements are associated with the options values estimated with an implicit convenience yield. Although such improvements are suggestive of potential gains in forecasting precision, more research is required before definitive conclusions can be reached about improving forecasts of basis convergence by including an options value variable.

When sufficient time has passed to allow observation of an adequately sized sample, this research can be reevaluated using price data on the basis of the new corn (or soybean) contract specification that was effective in the year 2000. The new contract provides for multiple delivery points along the Illinois Waterway, with varying discounts/premiums for the locations. A somewhat different timing option also remains in the contract. The USDA reports cash prices for many of these delivery sites, and these prices are likely more closely related to actual transactions than are those for the Chicago location. If so, the value of the new location options can be better estimated. We believe that our research demonstrates that it would be worthwhile to repeat the analysis if higher quality observations are available for cash prices.

The results in this article suggest that it is possible to estimate the joint values of implicit options in futures contracts and that these values appear to explain some of the variability of the basis at contract maturity. On the other hand, at least for our sample of data, this information had little effect on the precision of forecasts of basis convergence, that is, forecasts using this variable often did not reduce basis risk, and when it did, the reduction was small. Additional research can likely build on and refine the models used in this article, but the availability and quality of spot prices for commodities in alternative delivery locations (and for different grades) are potential limitations for such research.

APPENDIX

Estimated Joint Options Values, Without and With Convenience Yield[a]

		Without Convenience Yield								
		Year								
Month	*Ave.*	*89*	*90*	*91*	*92*	*93*	*94*	*95*	*96*	*97*
		(Cents per bushel)								
Mar	2.48	3.56	1.79	3.25	2.44	3.11	2.63	3.21	1.29	1.02
May	2.59	2.50	3.13	3.32	2.85	2.90	3.03	2.76	0.70	2.13
July	1.42	0.54	0.45	2.33	1.83	2.39	1.70	2.17	0.80	0.56
Sept	1.58	0.39	0.74	2.49	1.84	2.69	3.06	1.00	0.48	1.49
Dec	2.59	1.72	3.22	2.33	2.68	3.06	2.91	2.71	1.48	3.17

		With Convenience Yield								
		Year								
Month	*Ave.*	*89*	*90*	*91*	*92*	*93*	*94*	*95*	*96*	*97*
		(Cents per bushel)								
Mar	41.70	56.95	1.79	20.37	22.08	142.22	43.86	55.65	31.26	1.12
May	42.54	21.86	48.62	34.53	2.85	90.43	31.71	37.41	103.20	12.25
July	46.61	0.79	12.26	44.86	1.83	114.59	29.35	21.68	193.55	0.54
Sept	31.41	0.23	3.30	28.28	1.84	67.81	141.22	67.20	1.36	1.49
Dec	78.94	1.72	100.43	37.22	94.02	82.50	169.14	79.58	1.40	144.49

[a]Estimates based on Chicago and Toledo cash prices for the first delivery day of the month.

BIBLIOGRAPHY

Boyle, P. P. (1989). The quality option and timing option in futures contracts. The Journal of Finance, 44, 101–113.

Greene, W. H. (2000). Econometric analysis (4th ed.). Upper Saddle River, NJ: Prentice Hall.

Heifner, R. G. (1966). The gains from basing grain storage decisions on cash-future spread. The Journal of Farm Economics (now the American Journal of Agricultural Economics), 48, 1490–1495.

Hemler, M. L. (1988). The quality option in treasury bond futures contracts. Unpublished doctoral dissertation, University of Chicago.

Hranaiova, J. (2000). Delivery options in futures contracts and basis behavior. Unpublished doctoral dissertation, Cornell University, Ithaca, NY.

Hranaiova, J., Jarrow, R. A., & Tomek, W. G. (2001). Estimating the value of delivery options in futures contracts (Chicago Board of Trade Educational Research Foundation Working Paper 23B). Chicago: Chicago Board of Trade Educational Research Foundation.

Hranaiova, J., & Tomek, W. G. (2001). Delivery options in basis convergence (Chicago Board of Trade Educational Research Foundation Working Paper 23A). Chicago: Chicago Board of Trade Educational Research Foundation.

Jarrow, R. A., & Turnbull, S. M. (1996). Derivative securities. Cincinnati, OH: South-Western College Publishing.

Leuthold, R. M. (1979). An analysis of the futures-cash price basis for live cattle. North Central Journal of Agricultural Economics (now the Review of Agricultural Economics), 1, 47–52.

Liu, S., Brorsen, B. W., Oellermann, C. M., & Farris, P. L. (1994). Forecasting the nearby basis of live cattle. The Journal of Futures Markets, 14, 259–273.

Peck, A., & Williams, J. C. (1991). An evaluation of the performance of the Chicago Board of Trade wheat, corn, and soybean futures contracts during delivery periods from 1964–65 through 1988–89. A report to the National Grain and Feed Association (see also Food Research Institute Studies, 22, No. 2).

Pirrong, S. C., Kormendi, R., & Meguire, P. (1994). Multiple delivery points, pricing dynamics, and hedging effectiveness in futures markets for spatial commodities. The Journal of Futures Markets, 14, 545–573.

Silk, R. D. (1988). Implicit delivery options in futures contracts and optimal exercise strategy. Unpublished doctoral dissertation, Stanford University, Palo Alto, CA.

Working, H. (1953). Hedging reconsidered. Journal of Farm Economics (now the American Journal of Agricultural Economics), 35, 544–561.