Capital Asset Valuation Under Risk Aversion

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Stock prices are extremely volatile but dividends are not. This empirical observation poses an interesting puzzle for the conventional neoclassical theory of capital asset valuation which equates the price of a capital asset to the discounted sum of expected future dividends (Shiller). To highlight this point, some notation will be introduced. Denote the exdividend price of the i th capital asset at time t by P_{it}; D_{it} denotes the infinite stream of dividends for the i th asset. $0 < \beta < 1$ is the subjective rate of time preference and E_t represents conditional expectations based on all information available at time t. When capital markets are in equilibrium, assets are priced according to the familiar valuation formula,

$$(1) \quad P_{it} = \Sigma_j \beta^j E_t D_{it+j} .$$

Assuming that β is a constant, equation (1) implies that dividends are the only source of variation in capital asset values. Since this is clearly inconsistent with the evidence, at least from stock markets (LeRoy and Porter), formula (1) must be appropriately modified or discarded entirely. The search for a valuation principle more closely aligned with stylized facts of stock market behavior has been the focus of recent research in financial economics.

A fundamental modification follows when the assumption of risk-averse investors in capital markets is introduced (Lucas; Merton; Breeden). This modification introduces marginal rates of substitution in consumption between time periods into (1). More important, especially from the standpoint of the puzzle posed earlier, dividends are no longer the singular determinant of asset prices since consumption fluctuations influence asset prices as well. To make matters concrete, denote consumption at time t+j by C_{t+j} for j=0,1,2,...., . Then the modified asset valuation formula under the maintained hypothesis of risk aversion is,

$$(2) \quad P_{it} = \Sigma_j \beta^j E_t [\{U'(C_{t+j})/U'(C_t)\} D_{it+j}] ,$$

where $U'(C_{t+j})$ is the first-derivative of a von Neumann Mongenstern utility function evaluated at time t+j. U(.) is assumed to obey standard regularity conditions, namely, U(0)=0, U'(.) >0, and U''(.) <0. The first restriction says that no utility is obtained when nothing is consumed. Positive marginal utility of consumption and risk aversion are implied by the

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remaining restrictions. Equation (2) collapses to the conventional pricing formula (1) when risk neutrality is maintained . To confirm this result, consider a linear utility function $U(C_t)=\tilde{a}C_t$ with $U'(C_t)= U'(C_{t+1})=.....= \tilde{a}$. For this utility function, the marginal rate of substitution in consumption between any two time periods t and t+j, $U'(C_{t+j})/U'(C_t)$, is identically 1, whence the result. Clearly, the linear utility function violates the concavity restriction.

The economic intuition underlying the role of risk aversion in valuing capital assets merits further discussion. Consider the simplest case of a capital asset yielding an infinite stream of risky dividends with an expected value of $1 per time period. The representative consumer expects to consume $ C in current and all future time periods. Suppose now that the asset continues to yield an expected dividend of $1 per time period, although consumption in all future time periods is $ C'. Under these circumstances, a risk neutral investor places the same value on the capital asset as before; however, risk averse investors will likely revise their valuation of the capital asset. Changing the consumption stream has the effect of altering the marginal rate of substitution in all future time periods. Fluctuations in aggregate consumption alter asset values even when no change occurs in the character of the dividend series. The behavioral assumption of risk aversion, embodied in the non-linearity of the utility function, has significant implications to pricing assets.

The plan of this paper is as follows. Section 1 discusses the applicability of capital asset valuation models to farmland markets. The following section is devoted to a discussion of two estimation methods for valuation models. Direct estimation of closed-form solutions such as (2) poses formidable difficulties even when risk neutrality is imposed. A two-step instrumental variable estimation procedure is described. The second method relies on specification of a parametric form governing the evolution of future expectations. While both methods are feasible, section 3 presents an empirical application of the latter to aggregate land price and returns data. Closing remarks are in Section 4.

Valuation of Capital Assets

The literature on pricing agricultural land has an interesting and colorful history. Existing theories on farmland pricing have explained price changes by market returns (Melichar; Phipps), market returns and <u>ex post</u> capital gains (Castle and Hoch), farmland quality variations (Peterson), non-agricultural uses of farmland (Pope), variations in the inflation rate (Feldstein), speculative forces (Plaxico), and some combination of these factors including taxes (Robison et al). While this list is by no means exhaustive, it indicates the range of scholarly opinion on the subject. Ultimately, factors such as *quality differentials influence expectations of future profitability and are capitalized into land values. If farmland markets are efficient in Fama's sense of the term, then all available information is embodied in land values. Many analysts of land pricing adopt formula (1) which was shown to be a special case of formula (2) earlier. However, these studies do not make explicit the important fact that* ***these formulae are not ad hoc but follow as a logical implication of equilibrium in capital markets.*** *At least for this reason, formula (2) serves as a useful point of departure. Notice that the parametric structure of the valuation formula will depend upon assumptions made about the stochastic processes governing dividend and consumption growth. Another pertinent factor affecting the valuation formula is the form of the utility function. The sensitivity of (2) to varying assumptions illustrates these points.*

Dividend growth

Suppose that the stochastic process governing the evolution of dividends is a random walk, $[1-B]\ D_{it} = U_{it}$ *with initial value* $D_{it-1} = D_i^*$. U_{it} *is a white noise error term. For this model, the conditional expectation* $E_t D_{t+j}$ *is replaced by the minimum mean squared error forecast of* D_{t+j} *formed with information available at time t. That is, all future dividends are expected to remain at* Di^**, say, the long-run equilibrium value. Replacing conditional expectations in (1) by their corresponding optimal forecasts results in the valuation formula,*

$$(3) \quad P_{it} = \gamma D_i^*,$$

where the parameter $\gamma = \Sigma_j \beta_j = [1-\beta]^{-1}$. *Recent land pricing studies have either utilized (3) or a variant thereof after suitable modification of the discount rate (Alston; Burt; Pope, among others). According to (3), when capital markets are in equilibrium and dividends obey a random walk, the price of the i th capital asset is proportional to its dividend. The proportionality factor is intimately related to the discount factor. One common parametric form for* β *is* $[1+r]^{-1}$ *which amounts to setting* $\gamma = r^{-1}$ *(Burt).*

Now consider the Williams-Gordon-Rubinstein (WGR) dividend growth model. While this model was proposed originally for financial markets, land valuation studies have employed it in their analyses as well(Melichar; Pope, among others). In compact notation, the WGR model is $[1-\emptyset B]\ D_{it} = U_{it}$ *,* $abs[\emptyset] < 1$*, and initial value* $D_{it-1} = D_i^*$. *An AR(1) dividend growth process is implied by this model.* $\emptyset$ *is the mean dividend growth rate, assumed to be positive here. The random walk model discussed earlier results from setting* $\emptyset=1$ *in the WGR model. Following the same procedure, conditional expectations in (1) are replaced by minimum mean-squared error forecasts to obtain the appropriate pricing formula. In this case* $E_t D_{t+j} = \emptyset^j D_i^*$ *and , by substitution,*

(4) $Pit = \gamma^* D_i^*$.

An important difference belies formulae (3) and (4), namely that $= ^*$ only when $\emptyset=1$. Since D_i^* is independent of the index j, (1) implies that $\gamma^* = \Sigma_j[\beta\emptyset]^j = [1-\beta\emptyset]^{-1}$, given that $0 < \beta < 1$, abs($\emptyset$) < 1 , and $\emptyset > 0$. Notice that γ^* is now a function of $\emptyset$ so that non-constancy of cash value-land ratios can be attributed to changes in $\emptyset$. This concludes the discussion of dividend growth and its effect on valuation formulae.

Parametric Structure of the Utility Function

The most commonly employed parametric form in asset valuation studies is the constant relative risk aversion (CRRA) utility function (Grossman and Shiller; Litzenberger and Ronn). This specification has useful aggregation properties (Rubinstein) and includes other utility functions as nested alternatives. Empirical support for the CRRA hypothesis is well documented (Blume and Friend). Obvious similarities between the CRRA utility function and the Box-Cox transformation studied by analysts of risk attitudes in agricultural markets (Fleisher and Robison) can be identified. The CRRA utility function takes the form $U(C_t) = [1-\theta]^{-1} (C_t)^{[1-\theta]}$, with $0 < \theta < \infty$. The parameter θ is the coefficient of relative risk aversion. Under the maintained hypothesis of risk neutrality, $\theta = 0$ and $U(C_t) = C_t$, which is like the linear utility case discussed earlier. Testing the null hypothesis that $\theta=0$ assumes added significance because virtually all land pricing studies implicitly impose this restriction. Exploiting the property that $\lim U(C_t) = \ln C_t$ as the parameter θ approaches 1 yields the logarithmic utility function as a special case of CRRA. The logarithmic utility function exhibits decreasing absolute risk aversion in addition to CRRA. The CRRA function is parametrically parsimonious with only one parameter. This is a useful property from the perspective of statistical estimation involving time series with a limited number of observations. Other parametric forms of the utility function such as the quadratic could be employed as well.

Which specification of the utility function is employed affects the parametric form of marginal rates of substitution and, by implication, the valuation formula (2). For the CRRA utility function, the valuation formula (2) becomes,

(5) $P_{it} = \Sigma_j \beta^j E_t(C_{t+j}/C_t)^{-\theta} D_{it+j}$.

To explore this matter a little further, a straightforward argument establishes that the pricing formula associated with the logarithmic utility function simplifies to,

(5') $P_{it} = C_t \Sigma_j \beta^j E_t(D_{it+j}/C_{t+j})$,

given that consumption C_t is observed at time t and hence is nonstochastic. Asset prices here are proportional to the discounted sum of mean dividend-consumption ratios over an *infinite horizon. From comparing formulas (1), (5), and (5'), it is clear that assumptions made about risk attitudes are critically important to the empirical analysis of asset prices.*

Estimation of Asset Valuation Models

Before discussing estimation of model (2), a stronger justification for applying this methodology is needed. After all, model (2) emerged from a failure to reconcile volatile stock prices with sticky dividends (LeRoy and Porter; Shiller) and a similar justification may be lacking for land prices. The early literature on reconciling increasing real land prices with stable farm income becomes relevant here (Klinefelter; Tweeten and Martin). Model (2) easily resolves this paradox. Market returns to land have been extremely volatile in recent years unlike the era of stable income. Whether this volatility is consistent with prices implied by risk neutrality is essentially an empirical question which can be answered by taking recourse to variance bounds tests. The exponentially growing literature on these tests (Flavin; Marsh and Merton) preempts any complete discussion of this topic here. A further justification follows from considering reported non-constancy of cash value-land rent ratios (Robison et al). This phenomenon can be explained by consumption fluctuations. Econometric studies based on model (2) are justified at least by their rich conceptual foundations and also, as a secondary reason, because the commonly imposed risk neutrality assumption is a testable hypothesis.

Suppose one takes as given that models such as (2) more accurately replicate the true process generating, say, land price data. A natural direction to pursue then is development of procedures for estimation of risk aversion parameters, expectation formation parameters, etc.. Several methods have been proposed to achieve this goal. To motivate estimation issues, a useful point of departure is the observation that (2) is the solution to the stochastic euler equation ,

$$(6)\quad P_{it}\, U'(C_t) = \beta E_t U'(C_{t+1})(P_{it+1} + D_{it+1}).$$

Note that equation (6), which is a first-order stochastic difference equation, can be compactly written as ,

$$(6')\quad (1-\beta^{-1}B)E_t P_{it+1}\, U'(C_{t+1}) = -\beta^{-1}E_t U'(C_{t+1})D_{it+1}.$$

This is a special case of the second-order difference equation studied in the rational expectations literature on dynamic input demands (Hansen and Sargent). Because $0 < \beta < 1$, $\beta^{-1} > 1$ is the unstable root of (6'). Hence only the forward inverse of (6') can be evaluated. Solving (6') by this method gives (2), the forward-looking solution or, alternatively, iterating on expectations gives the result (Grossman and Shiller). Based on the equivalence of (3) and (6), econometric estimation of either form can be attempted.

Estimation by Instrumental Variables (Hansen and Singleton)

Suppose that the utility function is of the CRRA type. Construct new variables $Z_t=[C_{t+1}/C_t]^{-\theta}$ and $Y_t=(P_{it+1}+D_{it+1})/P_{it}$. Equation (6) can be rewritten as $E_t\beta Z_t Y_t - 1 = E_t f_t(\theta,\beta) = 0$ or, more compactly,

$$(7) \quad f_t(\theta,\beta) = \zeta_t ,$$

with the random variable ζ_t being white noise. $f_t(.)$ depends upon both the data and parameters. Consistent estimates of θ and β are obtained in two steps. In the first step, one obtains an estimate of a symmetric positive definite weighting matrix W which is naturally determined by the given sample of T observations augmented with a judiciously selected matrix of instruments (e.g., lagged values of consumption and land prices). Once W is determined, estimates of the parameters β and θ are obtained as a solution to the problem of minimizing the quadratic,

$$(8) \quad Q = g(\beta,\theta)'\, W\, g(\beta,\theta) ,$$

with respect to β and θ, where $g(\beta,\theta) = T^{-1} \Sigma_t f_t(\beta,\theta)$.

Estimation of the Closed-Form Solution (Litzenberger and Ronn)

The instrumental variables method involves direct estimation of the euler equation (6). As an alternative approach, estimation of the closed-form solution, say (5), can be attempted. Suitable assumptions about the parametric structure governing the evolution of future expectations yield such a closed-form solution. To illustrate this estimation strategy, suppose that expectations of dividends weighted by marginal rates of substitution evolve according to the second-order difference equation,

$$(9) \quad [(C_{t+1}/C_t)^{-\theta} D_{t+1} - (C_t/C_{t-1})^{-\theta} D_t] = \bar{a}_1 [(C_t/C_{t-1})^{-\theta} D_t - (C_{t-1}/C_{t-2})^{-\theta} D_{t-1}] + \xi_{1t}.$$

Equation (9) says that revisions in expectations in every time period are proportional to previous changes in dividends weighted by marginal rates of substitution. Under this forecasting scheme, the closed form-solution of (5) takes the specific form,

$$(10) \quad (P_t/P_{t-1}) = c_1 + c_2 (C_t/C_{t-1})^{-\theta} (D_t/P_{t-1}) + c_3\bar{a}_1 [(C_t/C_{t-1})^{-\theta}(D_t/P_{t-1}) - (C_{t-1}/C_{t-2})^{-\theta}(D_{t-1}/P_{t-1})] + c_3 (D_t/P_{t-1}) + \xi_{2t} ,$$

subject to certain non-linear parametric restrictions (Litzenberger and Ronn). Together equations (9) and (10) can be viewed as a system of nonlinear equations and estimated by a suitable procedure.

Several features of this system are noteworthy. First, parametric restrictions across equations characterize the structural model. Second, the maintained model reflects the aggregate stochastic behavior of consumption, prices, and dividends when capital markets are in equilibrium. In this sense, the model is not ad hoc and is a derivable implication of economic theory. Finally, equation (9) is in implicit function form so that it is impossible to disentangle endogenous variables explicitly in terms of exogenous variables.

An Application to Pricing U.S. Farmland

Data on farm real estate prices, returns to farmland investment, and aggregate farm consumption over the 1953-83 time period are needed to estimate models (9) and (10). Some of these data series are readily available. Others are not. Time series utilized in estimation and their limitations are discussed in the Appendix. Before implementing the estimation procedure, some assumptions about the random error terms were necessary. First-order serial correlation in each error term was assumed to prevail. Accordingly, the Durbin transformation was applied to each equation. The iterated 3SLS algorithm implemented on SAS was invoked in estimation. This method has an invariance property yielding identical estimates for equivalent model formulations, which is desirable for shared parameter systems such as the one estimated in this study.

Coefficient Estimates and Coefficient Stability

Parameter estimates for the maintained model are placed in Table 1. These estimates are with one exception statistically significant, at least at the 10% significance level. When a more stringent criterion (1%) is applied, four out of seven parameters pass the test. Empirical estimates support first-order serial correlation in the land pricing equation although a similar result for the dividend dynamics equation cannot be accepted.

The parameter θ is the coefficient of relative risk aversion. This coefficient is a measure of concavity of the utility function or, equivalently, the disutility of consumption fluctuations. A value of unity for θ would imply that land price movements are explained by a logarithmic utility function. Since the parameter θ takes a value of 4.66 with standard error 0.59, it is statistically significantly different from both zero and unity at the 1% level. By implication, the inference that neither linear utility nor logarithmic utility are consistent with aggregate U.S. agricultural data can be made.

The estimate of θ obtained in this study is consistent with the findings of other researchers. Litzenberger and Ronn obtained a value of 4.22 for their model of stock price variations. Friend and Blume obtain values ranging from 2.5 and 4.0 for this coefficient in their cross sectional model of stock price variations. Grossman and Shiller infer that a value of 4 for õ fitted the observed behavior of the S & P Index quite well. Another interesting exercise to perform is a study of coefficient stability. The data was subdivided into two time periods of equal length to determine whether the coefficients varied over time. A value of θ= 7.519 was obtained for 1953-67 and the corresponding value for 1968-83 was 4.437. These estimates were statistically significantly different at the 1% level. However, an F test to determine whether the data supported pooling of these two models supplied an answer in the affirmative.

Comparison With Risk Neutral Model

The empirical consequences of accounting for risk aversion are best judged by performing a comparative evaluation of model performance. Although the hypothesis of risk neutrality was unequivocally rejected, this model was estimated by setting õ=0 in both (9) and (10) for purposes of comparison. Estimated parameters for this version of the model are in Table 2. With only two statistically significant coefficients, it is evident that this model poorly explained the data.

Two measures of performance, root mean squared error (RMSE) and correlation between actual and predicted price changes (CORR), were computed. While RMSE was 0.0306 for the risk aversion model, a value of 0.0313 resulted for the risk neutral model. These did not appear to be numerically very different. CORR was 0.64 for the model maintaining risk aversion and 0.61 for the alternative model. Again, although the risk aversion model was marginally better, numerically similar values were obtained. Significant differences between models begin to show up for subsamples of the data set. During the 1968-83 period, the RMSE was 0.424 for the risk aversion model and 0.475 for the risk neutral model. This result favors the former, and is especially significant given that the 1968-83 historical period was turbulent for the aggregate U.S. farm sector. A further reinforcement regarding the superiority of the risk aversion model is obtained by a CORR value of 0.671 for the risk aversion model. The corresponding value was 0.521 for the alternative model. Hence actual land values are better correlated with predicted values for the model maintaining risk aversion.

Conclusions

An important omission in existing land valuation models is the effect of risk aversion. Accordingly, this paper attempted to motivate issues surrounding specification and estimation taking account of this omission. Preliminary results based on aggregate land price data support the empirical importance of accounting for risk aversion. However, these results are only tentative and more rigorous experimentation with alternative models is a priority item for future studies. Many questions need to be answered. First, variance bounds tests of the *consistency between agricultural asset price and returns variations must be performed. Second, the applicability of theoretical general equilibrium models used to justify model (2) should be subjected to closer scrutiny. While these models are appealing because the valuation formula derived is shown to be a property of equilibrium in capital markets, they may be inconsistent with structural peculiarities of agricultural markets. Third, the absence of aggregate farm consumption data makes it difficult to confirm or refute the maintained model without introducing some degree of arbitrariness. Finally, models based on risk aversion can be utilized in policy analysis. For example, questions such as the impact of tax reform on farmland prices can be answered. Because there are so many ways to improve on the themes addressed in this paper, the present effort may at best be viewed as a preliminary inquiry into the role of risk aversion in explaining asset price variation.*

Appendix: Data Used in Estimation

Data on farm real estate prices, returns to farmland investment, and aggregate farm consumption are needed for estimation. Value of farmland and buildings per acre are used as a measure of farm real estate price (USDA, 1984). Farm net income as a residual return to equity was divided by total acres of farmland to derive dividends accruing to investors (USDA, 1985). Consumption data on the U.S. farm sector are simply not available. The following procedure was utilized to obtain a proxy measure. First, a consumption function for the U.S. economy was estimated by regressing per capita real consumption (C) on per capita real disposable income (I), real interest rates (R), and a trend variable (T). The estimated consumption function (t-values in parentheses) was,

$$\begin{array}{l} \ln C = 1.3150 + 0.8407 \ln I - 0.0386 \ln R + 0.0046\, T\,. \\ \qquad\quad (1.92) \quad\; (10.36) \qquad\; (-3.08) \qquad\; (2.35) \end{array}$$

Data were deflated using the implicit income deflator, with 1982 equal to one. This data was obtained from <u>Economic Report to the President</u> . Next, data on per capita disposable income farm and nonfarm sources were collected (USDA,1984). Per capita disposable income of the farm population was approximated by proportioning taxes paid by the farm population to income derived from farm sources on the basis of importance of farm income relative to total income. Finally, per capita consumption of the farm population was estimated by substituting this income measure alongwith other variables into the aggregate U.S. consumption function.

Table 1: Estimated Parameters of CRRA Model

Parameter	Estimates+		
	1953-67	1968-83	1953-83
θ	7.519***	4.437***	4.659***
	(5.97)	(5.82)	(7.83)
c_1	1.140***	0.893***	0.870***
	(28.83)	(12.78)	(13.90)
c_2	-2.577**	5.662**	6.531**
	(-2.81)	(2.23)	(2.61)
c_3	-0.677	-0.845	-1.011*
	(-0.96)	(-1.11)	(-2.03)
$\tilde{a}_1$	-0.161	-0.558**	-0.588***
	(-0.43)	(-2.87)	(-6.34)
Rho_1	-0.041	-0.283	0.649***
	(-0.15)	(0.49)	(3.54)
Rho_2	-0.371	-0.088	-0.088
	(-0.92)	(-0.22)	(-0.39)

+ Numbers in parentheses are t-values.

* Significant at the 0.10 level.

** Significant at the 0.05 level.

*** Significant at the 0.01 level.

Table 2: Estimates for the Risk Neutral Model (θ=0)

Parameter	Estimates (1953-83)+
c_1	0.988* (43.01)
c_2	14.248 (0.01)
c_3	-13.035 (-0.01)
$\tilde{a}_1$	-0.068 (-0.01)
Rho1	0.574* (3.22)
Rho2	-0.096 (-0.02)

+ Numbers in parentheses are t-values.

* Significant at the 0.01 level.

References

Alston, Julian M. "An Analysis of Growth of U.S. Farmland Prices, 1963-82." Amer. J. Agr. Econ. 68(1986): 1-9.

Blume, Marshall and I. Friend. " The Asset Structure of Individual Portfolios and Some Implications for Utility Functions." J. of Fin. 10(1975): 585-603.

Breeden, Douglas. " An Intertemporal Asset Pricing Model with Stochastic Consumption and Investment Opportunities." J. Fin. Econ. 7(1979): 265-96.

Brock, William A. "Asset Prices in a Production Economy." in The Economics of Information and Uncertainty ed. J. McCall, Univ. Chicago Press, 1982.

Burt, Oscar R. "Econometric Modeling of the Capitalization Formula for Farmland Prices." Amer. J. Agr. Econ. 68(1986): 10-26.

Castle, Emery N. and I. Hoch. " Farm Real Estate Price Components, 1920-78." Amer. J. Agr. Econ. 64(1982): 8-18.

Economic Report of the President. U.S. Government Printing Office, Washington D.C., 1986.

Feldstein, Martin. "Inflation, Portfolio Choice, and the Prices of Land and Corporate Stock." Amer. J. Agr. Econ. 62(1980): 910-16.

Flavin, Marjorie A. "Excess Volatility in the Financial Markets: A Reassessment of the Empirical Evidence." J. Polit. Econ. 91(1983): 929-956.

Fleisher, Beverly and L.J. Robison. "Applications of Decision Theory and the Measurement of Attitudes Towards Risk in Farm Management Research in Industrialized and Third World Settings." MSU Int. Dev. Pap. No. 6, 1985.

Friend, I. and M. Blume. "The Demand for Risky Assets." Amer. Econ. Rev. 65(1975): 900-22.

Grossman, Sanford J. and R. J. Shiller. "The Determinants of the Variability of Stock Market Prices." Amer. Econ. Rev. 71(1981): 222-27.

Hansen, Lars P. and T. J. Sargent. "Formulating and Estimating Dynamic Linear Rational Expectations Models." J. Econ. Dynamics and Control 2(1980): 7-46.

Hansen, Lars P. and K. J. Singleton. "Generalized Instrumental Variables Estimation of Nonlinear Rational Expectations Models." Econometrica 50(1982): 1269-86.

Klinefelter, D. A. "Factors Affecting Farmland Values in Illinois." Illinois. Agr. Econ. 13(1973): 27-33.

LeRoy, Stephen F. and R. D. Porter. "The Present-Value Relation: Tests Based on Implied Variance Bounds." Econometrica 49(1981): 555-574.

Litzenberger, Robert H. and E.I. Ronn. "A Utility-Based Model of Common Stock Price Movements." J.Fin. 41(1986): 67-91.

Lucas, Robert E. "Asset Prices in an Exchange Economy." Econometrica 46(1978): 1426-46.

Marsh, Terry A. and R. C. Merton "Dividend Variability and Variance Bounds Tests for the Rationality of Stock Market Prices." Amer. Econ. Rev. 76(1986): 483-498.

Melichar, Emanuel. "Capital Gains Versus Current Income in the Farming Sector." Amer. J. Agr. Econ. 61(1979): 1085-92.

Peterson, Willis. "Land Quality and Prices." Amer. J. Agr. Econ. 68(1986): 812-819.

Phipps, Tim. T. "Land Prices and Farm-Based Returns." Amer. J. Agr. Econ. 66(1984): 422-29.

Plaxico, J. S. "Implications of Divergences in Sources of Returns in Agriculture." Amer. J. Agr. Econ. 61(1979): 1098-1102.

Pope, Arden C. "Agricultural Productive and Consumptive Use Components of Rural Land Values in Texas." Amer. J. Agr. Econ. 67(1985): 81-86.

Robison, Lindon J., D. A. Lins, and R. Venkataraman "Cash Rents and Land Values in U.S. Agriculture." Amer. J. Agr. Econ. 67(1985): 794-805.

Rubinstein, Mark. "The Valuation of Uncertain Income Streams and the Pricing of Options." Bell J. Econ. & Mgmt. Sci. 7(1976): 407-25.

Shiller, Robert J. "Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends." Amer. Econ. Rev. 71(1981): 421-36.

Tweeten, Luther G. and J. E. Martin. " A Methodology for Predicting U.S. Farm Real Estate Price Variation." J. Farm. Econ. 48(1966): 378-393.

U.S. Dept. Agr., ERS. Economic Indicators of the Farm Sector. U.S. Government Printing Office, Washington D.C., 1984.

U.S. Dept. Agr., ERS. Farm Real Estate: Historical Series Data, 1950-85. Stat. Bull. No. 738, Washington D.C., 1985.