AgEcon SEARCH
RESEARCH IN AGRICULTURAL & APPLIED ECONOMICS

The World's Largest Open Access Agricultural & Applied Economics Digital Library

This document is discoverable and free to researchers across the globe due to the work of AgEcon Search.

Help ensure our sustainability.

Give to AgEcon Search

AgEcon Search
http://ageconsearch.umn.edu
aesearch@umn.edu

*Papers downloaded from **AgEcon Search** may be used for non-commercial purposes and personal study only. No other use, including posting to another Internet site, is permitted without permission from the copyright owner (not AgEcon Search), or as allowed under the provisions of Fair Use, U.S. Copyright Act, Title 17 U.S.C.*

An Evaluation of Econometric Models of U.S. Farmland Prices

Rulon D. Pope, Randall A. Kramer, Richard D. Green, and B. Delworth Gardner

Previously published empirical models of U.S. farmland prices are reviewed and reestimated including recent data. It is apparent that structural changes have occurred. A simple single equation econometric model with less economic structure appears to forecast better than a simultaneous equation model. Finally, Box-Jenkins forecasts are roughly as good as those based upon a simultaneous equation econometric model, but somewhat inferior to the single equation model. The results suggest that further research may be needed to explain recent movements of farmland prices.

During the post World War II era, there has been much concern over rapidly rising agricultural land prices. Research by agricultural economists on the determinants of farmland prices appears to have peaked during the 1960's [Brake and Melichar]. However, the recent escalation in the rate of farmland price increases has caused a resurgence of interest [Morris, Harris, Gardner].

Though concern has been expressed over the ability of farm income to support farmland prices, there has been little effort devoted to the evaluation of econometric models of these prices in light of recent experiences. Clearly, policy decisions regarding commodity programs, financial instruments, and the impact of macroeconomic forces on farmland values and the distribution of wealth would benefit from a clear understanding of the farmland market. The concern of this paper is whether previously published models of the farmland market retain their structural credibility when estimated with recent data. The approach followed is not to critically evaluate or revamp these earlier models, but to examine their plausibility as explanations of recent market events, and study their predictive ability. After analyzing the reestimated models, one simultaneous equation and one single equation model are used in the evaluation of forecasting performance. To facilitate the analysis, a naive forecasting model (Box-Jenkins) without economic structure is compared with the econometric models. The naive forecasts are considered as benchmark results when evaluating the performance of the econometric models.

Some Econometric Models of Farmland Prices

Several simultaneous equation models of the U.S. farm real estate market have been developed. Three of the best known models are those presented by Reynolds and Timmons, Tweeten and Martin, and Herdt and Cochrane. All of the models did a reasonable job of explaining variations in land prices during the period for which they were originally estimated. To determine how well the models might perform now, they were reestimated utilizing more recent data. These

---

The authors are Assistant Professor, Post Graduate Research Associate, Assistant Professor and Professor, respectively, Department of Agricultural Economics, University of California, Davis.

Giannini Foundation Research Paper No. 524

The authors gratefully acknowledge helpful comments by the Editor and anonymous reviewers.

models are briefly reviewed and the results of the reestimation are discussed below.

Reynolds and Timmons used a two-equation recursive model for identifying the principal determinants of agricultural land prices for the period 1933-1965. They found that much of the variation in land prices could be explained by expected capital gains, predicted voluntary transfers of farmland, government payments for land diversion, conservation payments, farm enlargement, and the rate of return on common stock. When the model was reestimated with more recent data, 1946-1972, there were a number of changes in the signs and magnitudes of the coefficients. Table A1 in the Appendix details the definitions of variables used and the regression results. In the price equation four of the eight signs reversed and only one coefficient is statistically different from zero at the five percent level.

Tweeten and Martin presented a five-equation model for explaining changes in farmland values over time using recursive and ordinary least squares.[1] They found that the two major determinants of farm real estate price increases between 1950 and 1963 were capitalized benefits from government programs tied to land and pressures for farm enlargement. The model has been reestimated for the more recent period (1946-1972). Variable definitions and presentation of these regression results are found in Table A2 in the Appendix. Again, there was an abundance of sign changes and lack of statistical significance. For example, regardless of the estimation technique, all coefficients except lagged price (and possibly farm numbers) are statistically insignificant in the price equation.

The final simultaneous equation model considered is one presented by Herdt and Cochrane. They concluded that technological progress in conjunction with government supported output prices led to rising farmland prices. As with earlier models, this model also encounters problems when estimated with more recent data. However, generally sign reversals were few in number and lagged dependent variables are not used. For these reasons, this model was chosen as a representative of the simultaneous equation approach to be analyzed in more detail and utilized in forecasting later on.[2] Model formulation and estimation results are discussed in greater detail below.

The model is defined by

$$Ns = f(P, R, U, Lf)$$
(supply equation)

$$Nd = f(P, R, T, Pr/Pp, G)$$
(demand equation)

$$Ns = Nd$$
(identity)

where Ns is the number of farms supplied;[3] Nd is the number of farms demanded; P is the average value per acre of U.S. agricultural real estate in current dollars; R is the rate of return on nonfarm investment; U is the unemployment rate; Lf is the amount of land in farms; T is the USDA productivity index; Pr/Pp is the ratio of the index of prices received by farmers to the index of prices paid by farmers; and G is the wholesale price index.

Herdt and Cochrane estimated this model for 1913-1962 using two-stage least squares (2SLS). The model has been reestimated for the post-war years 1946-1972. In addition, the model has been estimated for 1913-1972. Since, *a priori* there is no reason to assume an absence of correlation across equations, both 2SLS and three-stage least squares (3SLS) estimates for the two time periods, as well as the Herdt and Cochrane original estimates, are presented in Table 1.

[1]Tweeten and Martin also employed a correction for autocorrelation for ordinary least squares. However, they discarded the model with autocorrelation when examining forecasting performance.

[2]A major reason for distinguishing the Herdt and Cochrane model from the other simultaneous equation models is its dissimilarity to Klinefelter's single equation model to be discussed later.

[3]Due to data limitations, farmland sold can only be characterized by the number of transfers, rather than acreage sold.

There is one sign change in each of the four sets of new estimates. In two of the sets of estimates, the sign of the coefficient for number of transfers in the demand equation reversed. This sign change may not be particularly meaningful since these estimates, and the original estimate of Herdt and Cochrane, are not significantly different from zero at conventional levels of type I errors. Also, the sign of the coefficient for the wholesale price index changed. Of the four new estimates for this coefficient, two are positive, two are negative, and all are statistically insignificant. Hence, the addition of ten more years of data makes it difficult to argue that this coefficient is nonzero.

In addition to these econometric models, a recent single equation model was tested which has less structural content than the other models, but fits the data well. Klinefelter assumes that the number of farm transfers is exogenous. Although the model is quite simple, the results have generated professional interest [Brake and Melichar]. Klinefelter found that 97 percent of the variation in Illinois land prices between 1951 and 1970 could be explained by net returns, average farm size, number of transfers, and expected capital gains. A model similar to Klinefelter's was estimated for U.S. data for the periods 1946-1972 and 1913-1972.[4] The results are presented in Table 2.

For the 1913-1972 estimates, the coefficients for net farm income and average farm size have unexpected signs. The coefficients for average farm size and number of transfers are not significantly different from zero. When the model was estimated for 1946-1972, the expected signs for net farm income and average farm size were obtained, but the sign for number of transfers reversed. Despite several implausible signs, this model was utilized for forecasting due to the high percentage of price variation explained by the variables ($\bar{R}^2$ of .952 (1913-1972) and .989 (1946-1972)) and for the reasons mentioned above.[5]

Forecast Results — Econometric Models

In order to forecast with the Herdt and Cochrane model, the reduced form equation for price was calculated and then the values of the exogenous variables were substituted to solve for price. Thus, the forecasts are ex post in the sense that actual values of the exogenous variables are used. The results are presented in Table 3. On the basis of root mean square error (RMSE), the various versions of the Herdt and Cochrane model can be compared. It is apparent that for within-sample forecasting, both sets of 2SLS estimates outperformed the 3SLS estimates. This is a rather surprising result since econometricians generally prefer 3SLS over 2SLS due to a presumption of the latter's lack of asymptotic efficiency. However, the better forecasting performance of the 2SLS estimates may result from the fact that full information estimation methods, such as 3SLS, are more sensitive to specification error than are the k-class estimators such as 2SLS. Since 3SLS takes into account the correlation between the disturbances of all the structural equations, a specification error in one equation will affect all of the coefficient estimates of the system.

None of the Herdt and Cochrane reduced form equations forecasted well beyond the sample. For example, the actual undeflated value of farm real estate per acre was \$340.48 in 1975. The highest forecast for that year was

[4]The model differed from Klinefelter's model as follows: (a) net farm income was used in place of net returns to landlords; (b) instead of deflating variables, the GNP deflator was entered as an explicit variable; (c) in the calculation of capital gains, capital improvements were subtracted out. These changes were required either to accommodate U.S. data or to achieve consistency with the other model used in forecasting.

[5]Of course a high $\bar{R}^2$ does not necessarily imply that a model will forecast well. Moreover, the Durbin-Watson statistic for the period 1946-1972 suggests evidence of positive autocorrelation. However, since expected capital gains contains transformations of lagged values of the dependent variable, the Durbin-Watson statistic may not be appropriate. If autocorrelation is present, parameters estimates will be inconsistent.

TABLE 1. Estimation Results for the Herdt and Cochrane Model[a]

	Original estimates	New estimates			
	2SLS 1913-1962	2SLS 1913-1972	2SLS 1946-1972	3SLS 1913-1972	3SLS 1946-1972
Supply Equation					
P	.064 (.538)	1.29 (6.123)	.244 (1.147)	1.106 (3.394)	.239 (1.318)
R	−5.672 (−4.634)	−19.42 (−7.162)	−1.33 (−.336)	−17.04 (−4.075)	−1.363 (−.405)
U	−.789 (−4.197)	−.337 (−2.132)	−.597 (−.542)	−.54 (−2.312)	−.892 (−.953)
Lf	.004 (1.333)	.036 (6.078)	.00001 (3.084)	.00003 (3.374)	.00001 (3.527)
Demand Equation					
Nd	−1.043 (−1.496)	−1.17 (−2.269)	.729[b] (.409)	−1.36 (−2.152)	.550[b] (.387)
R	8.315 (3.795)	18.94 (16.798)	16.38 (2.622)	19.90 (13.994)	17.166 (3.438)
T	1.699 (5.293)	2.35 (7.322)	2.22 (3.248)	2.64 (6.731)	2.296 (4.178)
Pr/Pp	.757 (2.035)	1.00 (2.804)	.3995 (.625)	1.22 (2.827)	.377 (.737)
G	.379 (2.399)	−.0335[b] (.205)	.669 (.547)	−.213[b] (−1.074)	.417 (.426)

[a]T-ratios are shown in parentheses.
[b]Denotes sign change as compared to original estimates.

TABLE 2. Estimation Results for the Modified Klinefelter Model[a]

Variables[b]	1913-1972	1946-1972
net farm income	−.0047 (−5.752)	.0036 (3.128)
average farm size	−.0536 (−0.683)	.5683 (7.0562)
number of transfers	−.0250 (−1.283)	.9526 (5.705)
expected capital gains	2.4099 (4.131)	.2203 (0.575)
GNP deflator	2.6843 (7.694)	1.1363 (3.781)
$\bar{R}^2$	.952	.989
Durbin-Watson statistic	2.581	.706

[a]T-ratios are given in parentheses.
[b]The dependent variable is the average value of U.S. farm real estate per acre.

TABLE 3. Econometric Forecasts of Farmland Prices[a]

	1973	1974	1975	RMSE[b] within sample	RMSE beyond sample
Actual	238.14	297.80	340.48	--	--
Herdt Cochrane					
2SLS 1913-72	196.84	212.20	222.07	7.14	87.73
2SLS 1946-72	218.65	246.85	269.40	10.49	51.00
3SLS 1913-72	198.58	214.69	224.06	7.74	85.81
3SLS 1946-72	204.68	228.84	244.12	20.25	68.91
Modified Klinefelter					
1913-72	212.52	238.81	298.04	10.36	44.49
1946-72	224.44	257.72	284.88	4.73	40.35

[a]Forecasts of undeflated value of U.S. agricultural land and buildings per acre.
[b]Root mean square error.

\$269.40, and the lowest was \$222.07. On the basis of RMSE beyond the sample, the 2SLS 1946-1972 estimates performed the best, followed by the 3SLS, 1946-1972 estimates.

The forecast results for the modified Klinefelter model are also presented in Table 3. For within-sample forecasts, the 1946-1972 estimates did better than any of the Herdt and Cochrane reduced forms. For both time periods the Klinefelter model forecasted better beyond the sample than each of the Herdt and Cochrane versions. It is apparent from these results that a simple model with implausible signs (e.g., coefficient of transfers) can still forecast quite well.

In the modified Klinefelter model expected capital gains (a three year moving average) includes lagged values of the dependent variable price. One is curious whether time series models based solely on the lag structure of the dependent variable plus more general error structures might possess as great a predictive power as the economic models. In the following section, Box-Jenkins forecasts are presented and later compared with the econometric forecasts. The Box-Jenkin's results are viewed as benchmark forecasts, since it is generally hoped that econometric models perform at least as well as naive statistical models.

Forecast Results — Time Series Model

As an alternative to the econometric models, time series models of an integrated autoregressive moving average form are used to obtain forecasts of land prices [Box and Jenkins].[6] These are statistical models of the form

$$Z_t = \phi_1 Z_{t-1} + \ldots + \phi_p Z_{t-p} + \\ \delta + U_t - \Theta_1 U_{t-1} - \ldots - \Theta_q U_{t-q}$$

where the Z's are observations generated by a stochastic process, the U's are independently distributed random variables with mean zero and constant variance, and δ, ϕ_i, and Θ_i are unknown parameters.

The first part of the model is referred to as the autoregressive portion and the latter part as the moving average portion. The term autoregressive derives from the fact that the first part of the above model is essentially a regression equation in which Z_t is related to its own past values rather than a set of independent variables. The moving average term comes from the fact that the second portion of the time series model is just a moving average of the disturbances reaching back for q periods. If the observations are in difference form, then the process is called an integrated autoregressive-moving average process

[6]Agricultural economists are not as familiar with this technique as with econometric models; however, there have been some applications of this procedure in agricultural economics. See, for example, Oliveira and Rausser, and Schmitz and Watts.

(ARIMA). Differencing of the data is often necessary in order to convert the process into a stationary one; that is, to yield a time series in which the joint distribution of any subset of observations of the series remains unchanged when the same constant is added to the time subscript of each observation. A related concept is that of weakly stationary in which a time series mean and autocovariance function are independent of time. In most cases first or second differencing of the original series suffices to convert a nonstationary time series into a stationary one. From an estimation viewpoint, stationarity results in a reduction of the number of unknown parameters to be estimated.

The first stage in selecting an appropriate time series model is to properly identify the process generating the observations. This is done by examining the estimated autocorrelation and partial autocorrelation functions. Box and Jenkins (pp. 176-77) provide tables describing the nature of the theoretical autocorrelation functions for various ARIMA processes. In general, the identification procedure is based on the characteristic behavior of autocorrelations and partial autocorrelations for known ARIMA processes. For example, if the autocorrelations exhibit spikes at lags 1 through q, then cut off and the partial autocorrelations tail off, then the process is a qth order moving average process. Other characteristics describe autoregressive and mixed autoregressive-moving average processes. These characteristics, however, describe the behavior of theoretical autocorrelation and partial autocorrelation functions. To analyze the estimated functions, the standard errors are needed to determine if the spikes are significantly different from zero at various lags.

From the estimated autocorrelation and partial autocorrelation functions, given in Table 4, based on 1913-1972 observations, the model was identified as an ARI(2,2) or possibly an IMA(2,2); that is, an integrated autoregressive process of order 2,2 or an integrated moving average process of order 2,2. To see this, observe the patterns of the autocorrelation and partial autocorrelation functions in Table 4. For the original series the autocorrelations remain relatively large for several lags suggesting a nonstationary series. The standard errors associated with the first and second lags are 0.13 and 0.21 respectively. Thus, at least one differencing is necessary. For the first difference, only the autocorrelations of lags three and five are significantly different from zero. Their associated standard errors are 0.13 and 0.14. This may suggest some first order process, but based on graphical inspection of the first difference and that of the second difference together with the autocorrelations and partial autocorrelations, a higher order model appeared to be more appropriate.

The autocorrelations of the second difference are relatively large for the first two lags with standard errors of 0.13 and 0.14 respectively. However, there are some rather large values of the autocorrelations for larger lags suggesting a mixture of exponentials or damped sine waves. The same conclusions are reached by considering the values of the partial autocorrelations. Thus, these observations suggest an ARI(2,2) or possibly an IMA(2,2) process.

Based on these identifications, the models were estimated. The estimated results are:

$$W_t = 1.071\ W_{t-1} - 0.215\ W_{t-2} + U_t \qquad \text{ARI(2,2) model,}$$

$$W_t = U_t + 1.104\ U_{t-1} - 0.023\ U_{t-2} \qquad \text{IMA (2,2) model,}$$

where W_t represents the second difference of the Z_t.

Diagnostic checks were made on both of the estimated models. The estimated autocorrelations of the residuals were used to evaluate the goodness of fits of the models. If, for example, $W_t = 1.071\ W_{t-1} - 0.215\ W_{t-2} + U_t$ were approximately equal to the true model, then the estimated residuals, $\hat{U}_t$, would constitute a white noise process. Thus, the sample autocorrelations of the estimated residuals would be approximately uncorrelated. Any departure from small values of au-

TABLE 4. Autocorrelation and Partial Autocorrelation Functions for 1913-1972 Farmland Prices

	Autocorrelations									
	1[a]	2	3	4	5	6	7	8	9	10
Original series	0.91	0.86	0.79	0.70	0.64	0.57	0.51	0.46	0.40	0.35
	0.30	0.26	0.21	0.16	0.12	0.08	0.05	0.01	−0.02	−0.06
First differences	0.11	0.09	0.28	0.13	0.31	0.19	0.13	0.21	0.07	0.28
	0.08	−0.07	0.16	−0.08	0.08	0.06	0.01	0.06	−0.10	0.03
Second differences	−0.29	−0.20	0.10	−0.18	0.21	−0.05	−0.06	0.09	−0.20	0.27
	0.00	−0.22	0.16	−0.19	0.11	0.05	−0.03	0.10	−0.19	0.04
	Partial autocorrelations									
	1	2	3	4	5	6	7	8	9	10
Original series	0.91	0.13	−0.09	−0.13	0.01	0.01	−0.01	0.01	−0.04	−0.01
	−0.04	−0.02	−0.03	−0.10	0.04	0.01	−0.01	−0.04	−0.08	−0.04
First differences	0.11	0.07	0.27	0.08	0.28	0.09	0.07	0.06	−0.06	0.17
	−0.09	−0.17	−0.03	−0.22	0.02	−0.05	0.12	0.03	−0.04	0.06
Second differences	−0.29	−0.31	−0.08	−0.28	0.07	−0.07	−0.00	0.02	−0.15	0.20
	0.10	−0.04	0.06	−0.14	0.01	−0.08	0.10	0.04	−0.04	−0.05

[a]Represents the order of lag.

tocorrelations would indicate inadequacies in the fitted model. In both models, all of the autocorrelations were not significantly different from zero. In addition, the Box-Pierce Q statistic

$$Q = T \sum_{j=1}^{K} r_j^2$$

where r_j are the estimated autocorrelations of the residuals, T = number of W's used to fit the model, and Q is approximately chi-square distributed with (K-p-q) degrees of freedom, was calculated to determine the goodness of fit of the models.[7] Using K = 20, the number of lags, for both models the value of Q, 17.75 and 13.14, was small relative to the critical value of $\chi^2_{.05,18} = 28.9$. The values of the Box-Pierce Q statistic indicate that the whole set of sample autocorrelations for lags 1 through 20 taken as a whole are small. Thus, the residuals from both of the models, based on the Q statistic, tend to indicate that the models fit the observed data well.

[7]For a more detailed discussion of the use of the Q statistic see Box and Jenkins, pp. 290-291.

The forecasting performances of the above estimated models were examined by predicting land prices within and outside the sample. The within sample forecasts were based on one period forecasts for the period 1913-1972, and the forecasts outside the sample were for the years 1973, 1974 and 1975. The results are presented in Table 5. For comparative purposes, results from a logarithmic model for the years 1913-1972 are also presented. Though substantially reducing the degrees of freedom, the postwar years were also estimated separately because of the land price spiral during this period.

The empirical results show that all of the models performed much better within rather than outside of the sampling period. However, the outside forecasts are one, two, and three period ahead forecasts, whereas the within sample forecasts are all one period ahead forecasts. Furthermore, forecasts obtained from the estimated model based on 1946-1972 data appear superior to those of other models. However, the estimated standard errors of the coefficients were high due to the relatively small number of observations used. The rule of thumb in estimat-

TABLE 5. Box-Jenkins Forecasts of Farmland Prices

	1973	1974	1975	RMSE[a] within sample	RMSE beyond sample
Actual	238.14	297.80	340.48	--	--
IAR (2,2)	214.54	223.14	231.73	6.73	77.37
upper limits[b]	227.02	240.17	253.39		
lower limits	202.06	206.10	210.06		
IMA (2,2)	218.05	225.74	233.42	6.25	75.41
upper limits	229.88	241.61	251.98		
lower limits	206.23	209.86	214.85		
ARIMA (2,2,2)[c]	228.49	242.17	257.51	6.01	57.94
upper limits	264.08	313.55	362.12		
lower limits	197.69	187.04	183.12		
ARIMA (1,1,1)[d]	243.20	249.95	251.13	7.38	58.59
upper limits	259.49	293.52	314.32		
lower limits	226.91	206.38	188.40		

[a]Root mean square error.

[b]Upper and lower limits for 95 percent confidence intervals.

[c]Model based on data in logarithmic form.

[d]Model based on 1946-1972 sample period.

ing time series models is that at least 50 observations are needed to adequately estimate a model [Box and Jenkins, p. 18].

The logarithmic model performs relatively well. On the basis of RMSE, it outperforms all other time series models including the postwar model. The implication of the logarithmic model is that percentage changes (rather than the level of changes) have remained relatively stable through time.

A Brief Comparison of Econometric and Time Series Forecasts

The simultaneous equation econometric model used in this study yielded forecasts about as accurate as the benchmark forecasts of the Box-Jenkins method for the postwar years when land prices were rapidly escalating. For the longer time period (1913-1972), the time series models performed better than the simultaneous equation model on the basis of RMSE, both within and beyond sample. For this same period the Klinefelter model had the lowest beyond-sample RMSE. Further, the Klinefelter model performed better than either time series or the simultaneous equation econometric model for the postwar years. Overall, the poorest predictors appear to be generated by the simultaneous equation models.

Conclusions

It is not uncommon when comparing time series and econometric forecasts to discover that time series models provide as good or better short-term forecasts than econometric models.[8] The above results are consistent with this conclusion. However, the single equation model predicted well too, and it may generate the best predictors. This result is surprising, particularly since the magnitudes and signs of the coefficient estimates appear very sensitive to the sample period used. Also, although the model may have microeconomic foundations, it explains little as a market model. Since expected capital gains are functionally related to lagged land prices, it appears that more study is needed to explain the recent rise in farm prices and capital gains.

The simultaneous equation models presumably have greater causal foundations reflecting market behavior of sellers and buyers. However, results in Tables 1, A1 and A2, indicate that attempts to incorporate greater structural detail in econometric models of the land market have not instilled much confidence in their structural performance. One would expect the magnitude of parameter estimates to be sensitive to the sample period. However, when recent data were added to the sample numerous changes in signs of coefficients occurred for all of the simultaneous equation models. Further, most of the estimated coefficients were not statistically significant from zero.

These results suggest that the model specifications do not reflect accurately enough the relevant structural changes and other characteristics of the farmland market. Therefore, if one is concerned with both predictive ability and economic structure, additional research is needed to explain recent movements of farmland prices. Such research may seek a better understanding of how expectations are formulated by land market participants, and a better understanding of the motives for holding real versus liquid capital in an inflationary economy. Further effort directed towards a set of more general statistical assumptions may also prove fruitful. For example, one could integrate the ARIMA Box-Jenkins models and econometric models as recently suggested by Newbold and Davies. Given the interest in the farmland market, additional research should be beneficial to farmers and policy makers.

[8]For an interesting discussion of the relative merits of Box-Jenkins versus econometric models see Naylor, Seaks, and Wichern.

References

Box, G. E. P., and G. M. Jenkins. *Time Series Analysis Forecasting and Control.* San Francisco: Holden-Day, 1970.

Brake, J. R., and E. Melichar. "Agricultural Finance and Capital Markets," in L. R. Martin (ed.), *A Survey of Agricultural Economics Literature,* Volume 1, University of Minnesota Press, Minneapolis, 1977.

Gardner, B. D. "Issues Affecting the Availability and Price of Land for Agriculture." Paper presented, California Chapter Meeting of the American Society of Agronomy, Fresno, January 26, 1978.

Harris, D. G. "Inflation-Indexed Price Supports and Land Values." *American Journal of Agricultural Economics,* 59 (1977):489-495.

Herdt, R. W., and W. W. Cochrane. "Farmland Prices and Technological Advance." *Journal of Farm Economics,* 48 (1966):243-263.

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

Morris, D. "Farmland Values and Urbanization." *Agricultural Economics Research,* 30 (1978):44-47.

Naylor, T. H., T. G. Seaks, and D. W. Wichern. "Box-Jenkins Methods: An Alternative to Econometric Models." *International Statistical Review,* 40 (1972):123-137.

Newbold, P. and N. Davies. "Error Misspecification and Spurious Regressions." *International Economic Review,* 19 (1978):513-519.

Oliveira, R. A. and G. C. Rausser. "Daily Fluctuations in Campground Use: An Econometric Analysis." *American Journal of Agricultural Economics,* 59 (1977):283-293.

Reynolds, T. E., and J. F. Timmons. *Factors Affecting Farmland Values in the United States.* Iowa Agricultural Experiment Station Research Bulletin 566, 1969.

Schmitz, A. and D. G. Watts, "Forecasting Wheat Yields: An Application of Parametric Time Series Modeling." *American Journal of Agricultural Economics,* 52 (1970):247-262.

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

APPENDIX TABLE A1. Estimation Results for the Reynolds and Timmons Model[a]

	Price equation									
	Constant	T[c]	GPL[d]	CP[e]	CG[f]	CG(33-41)	A[g]	1/R[h]	NFI[i]	NFI (56-65)
Original estimates 1933-1965	72.99 (7.73)	−0.23 (−5.75)	11.97 (5.07)	7.55 (2.64)	2.37 (3.76)	−2.25 (−4.09)	0.38 (0.86)	0.73 (3.65)	2.02 (4.21)	0.23 (0.74)
New estimates 1946-1972	36.533	0.033[b] (0.717)	23.663 (6.396)	−96.322[b] −(1.786)	0.117 (0.443)		−0.001[b] (−0.020)	22.923 (1.265)	−2.885[b] (−1.018)	

	Transfers equation							
	Constant	Constant (42-47)	D/E[j]	LA[k]	E(F/NF)[l]	A	CG	CG(33-41)
Original estimates 1933-1965	−11.23 (−0.32)	31.67 (3.81)	−7.36 (−7.08)	13.76 (12.18)	1.58 (3.36)	−3.35 (−2.19)	4.86 (3.20)	−2.96 (1.91)
New estimates 1946-1972	−309.994		19.690[b] (2.997)	34.824 (8.640)	−7.700[b] (−0.080)	−0.510 (−1.336)	−1.496[b] (−0.412)	

[a]T-ratios are shown in parentheses.
[b]Denotes sign change as compared to original estimates.
[c]Transfers of farmland.
[d]Government payments for land diversion.
[e]Conservation payments to farmers.
[f]Moving average of capital gains on farmland.
[g]Change in average farm size.
[h]Inverse of the rate of return on nonfarm investment.
[i]Moving average of net farm income.
[j]Debt to equity ratio.
[k]Manhours of labor used in farming.
[l]Moving average of the ratio of farm to nonfarm earnings.

APPENDIX TABLE A2. Estimation Results for the Tweeten and Martin Model[a]

Land price equation

	Constant	L_t^c	T_t^d	A_t^e	F_{t-1}^f	r_{t-1}^g	P_{t-1}^h	R^2	d
Original estimates:									
OLS 1923-1963	89.25	−.037 (1.892)	−.23 (2.49)	−.0027 (1.164)	.58 (3.04)	−1.56 (2.38)	.77 (7.09)	.95	1.53
RLS 1923-1963	88.58	−.033 (1.906)	−.41 (4.06)	−.0011 (.5138)	.58 (3.23)	−1.63 (2.90)	.77 (8.83)		
New estimates:									
OLS 1946-1972	14.623	.025[b] (.851)	.113[b] (.544)	−.009 (−2.358)	1.147 (1.893)	−1.942 (−1.206)	.889 (7.332)	.99	2.26
RLS 1946-1972	41.890	−0.011 (−0.302)	−0.298 (−1.078)	−0.004 (−1.011)	1.184 (1.679)	1.567 (−0.882)	0.938 (8.522)		

Land-in-farms equation

	Constant	C_t^i	F_{t-1}	LR_t^j	E_t^k	L_{t-1}	R^2	d
Original estimates:								
OLS 1923-1963	−134.37	.41 (2.20)	.80 (2.37)	.36 (1.96)	−.67 (2.75)	1.00 (38.31)	.99	.87
RLS 1923-1963	−83.21	.28 (.89)	.79 (2.21)	.25 (.93)	−.64 (2.49)	1.00 (34.19)		
New estimates:								
OLS 1946-1972	986.291[b]	0.149 (.219)	−7.113[b] (−1.730)	−0.481[b] (−0.817)	−1.379 (−3.345)	0.257 (1.636)	.66	1.99
RLS 1946-1972	1508.5[b]	−1.900 (−1.334)	0.652 (0.104)	−2.116[b] (−1.815)	−1.106 (−2.494)	0.314 (2.023)		

Cropland equation

	Constant	F_{t-1}	LR_t	T_2^l	C_{t-1}	R^2	d
Original estimates:							
OLS 1923-1963	180.73	0.38 (1.73)	−0.44 (5.92)	−3.13 (1.22)	0.51 (5.13)	.91	1.51
New estimates:							
OLS 1946-1972	111.156	2.890 (3.433)	−0.298 (−1.753)		0.560 (3.050)	.94	1.88

APPENDIX TABLE A2 (Continued)

	Farm numbers equation							
	Constant	JX_{t-1}[m]	CG_{t-1}[n]	S_t[o]	T2	A_{t-1}	R^2	d
Original estimates: OLS 1923-1963	418.46	0.11 (1.05)	−9.72 (4.86)	.017 (3.72)	49.92 (1.96)	0.94 (38.75)	.99	.90
New estimates: OLS 1946-1972	343.475	29.125 (1.385)	0.091[b] (0.071)	−0.035 (−0.868)		0.93 (21.346)	.98	2.81

	Farm transfers equation							
	Constant	JX_{t-1}	CG_{t-1}	S_t	T2	T_{t-1}	R^2	d
Original estimates: OLS 1923-1963	51.16	−0.028 (1.897)	−0.67 (3.16)	−0.0007 (2.132)	11.58 (4.08)	0.32 (2.14)	.87	1.72
New estimates: OLS 1946-1972	8.370	0.137[b] (0.367)	−0.033 (−1.425)	−0.0006 (−0.867)		0.839 (7.936)	.89	1.10

[a]T-ratios are shown in parentheses.
[b]Denotes sign change as compared to original estimates.
[c]Land in farms.
[d]Transfers of farmland.
[e]Number of farms.
[f]Net farm income.
[g]Rate of return on nonfarm investment.
[h]Price index of farm real estate deflated by the wholesale price index.
[i]Cropland used for crops.
[j]Land removed from production by government programs.
[k]Nonfarm employment.
[l]Dummy variable equal to 1 from 1942 to 1948.
[m]Ratio of nonfarm to farm earnings.
[n]Moving average of capital gains on farmland.
[o]Stock of machinery in constant dollars.