DETERMINANTS OF FARMLAND PRICES IN A DYNAMIC ERROR CORRECTION FORM: A NEW ZEALAND CASE

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DETERMINANTS OF FARMLAND PRICES IN A DYNAMIC ERROR CORRECTION FORM: A NEW ZEALAND CASE

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July 1999

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Determinants of Farmland Prices in a Dynamic Error Correction Form: 
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Abstract

This paper examines the effect of real net farm residual income, inflation, 
and the real interest rate on the movement of real farmland prices during 
1970 to 1997 using a parsimonious error correction model. The empirical 
evidence suggests that the long run trend in the real land price can be 
explained by real net farm residual income to land and the real interest rate, 
with inflation having no statistically significant effect. However, the growth 
in the real interest rate explains about two thirds of the growth in land prices 
in the immediate to short-run.

1. Introduction

The purpose of this study is to investigate the role of different economic variables that are 
hypothesised to explain the historical land price movement in the New Zealand context. 
Nominal real estate farmland prices in New Zealand have generally trended upward during 
1970 to 1997 (see Figure 1, Panel (a)). Up to 1983, the nominal price index exhibits a 
continuous increase, but from 1984 to 1989 it shows a downward trend. During 1990 and 
1991, nominal prices are almost stable and from 1992 they increased rapidly until stabilising 

As shown in Panel (b), the real price index exhibits two main episodes. The first episode 
(1970 to 1985) was characterised by a generally increasing real price level whereby the index 
increased from 34.18 in 1970 to a high of 478.42 in 1985. During the second episode the 
index decreased to 38.97 by 1997. Throughout the entire 28 year period, however, there was 
substantial volatility in the real price level. That is, during the episode of generally increasing 
prices there were several years in which the index actually decreased, whilst in the episode of 
decreasing prices the index increased in some years. Panel (c) shows that the volatility in the 
real price index was greater during the second episode, especially after 1992.
Figure 1. Average farmland price movements from 1970 through 1997.

(a) Nominal Land Price (NLP)

(b) Real Land Price (LP), 1989=100

(c) Percent Change in LP (DLLP)
In earlier studies, Nartea and Pellegrino (1997) and Nartea and Dhungana (1998) have pointed out that more than two thirds of the historical mean return (nominal) to farmland was contributed from capital gains. This increased capital gain is mainly due to the increase in the farmland value. Addressing the low farm returns gained in 1996 season, the National Bank of New Zealand (1997) reflects its concern in a recent report that ‘the rate of capital gain on rural land can not be fully justified by the rate of increase in gross profits’ and that ‘the value of rural land in relation to income in 1996 is perhaps as far out of balance as it has been for some time’. A very similar remark can be found in Peters (1966) describing the movement of land prices over time in England and Wales. Peters noted that during the period 1958-65, the annual yield from land appeared to be too low to form the basis for land values and remarked that land prices were sustained above their capitalised annual values by speculation, amenity considerations, hedging against inflation and taxation concessions. These remarks regarding the determination of farmland price movements have raised the interest to look into the role of economic variables driving the movement in the land prices in New Zealand Agriculture.

This paper examines the magnitude and causes of real estate farmland price movements, primarily by means of the error correction model based on co-integration analysis. As guided by micro-economic theory, land can be regarded as a productive asset with its value being determined by the capitalisation of its income. Besides farm income, this research investigates the effect of macro-economic and institutional factors on the movement of real estate farmland prices in New Zealand during the 1970 to 1997 period.

2. Research Method and Estimation Procedure

2.1 Error Correction Model

The error correction model as a dynamic specification has been gaining popularity in macro-econometric modelling following the work of Granger & Newbold (1977), and its use in Davidson, Hendry, Srba, and Yeo (1978). The model, as such, incorporates short-run as well as long-run information into the equation that comprises stationary components. It has additional advantages in that it can be estimated consistently using ordinary least squares (OLS) and performs well empirically (Engel and Granger, 1987). Importantly, it offers modelling of long-run as well as short-run relationships between integrated series since it contains variables in levels as well as in differences. The traditional approach of reformulating the equation in terms of differences rather than in levels to avoid the danger of spurious regression, as suggested by Granger and Newbold (1977); Nelson and Kang (1984), has received severe criticism from the econometricians since it looses the long run relationship that may exist between the variables of concern. Alternatively, the error correction model, which provides the short-run as well as the long run relationship between the variables, has been developed. The correspondence of the error correction model’s notion of a long run relationship to the statistical concept of co-integration has been explored by Engle and Granger (1987). Co-integration only provides the long run or equilibrium properties explained by economic theory. Indeed, economic theory itself usually has very little to say regarding the dynamic process by which variables move towards equilibrium (Lloyd and Rayner, 1990). Engle and Granger (1987) have concluded that if two or more series are all integrated to order 1, i.e. I(1), and are cointegrated then there exists an error correction mechanism that tends to eliminate part of each period’s long-run equilibrium error
for both short-run and long-run variables. For simplicity, an error correction form involving two I(1) variables, y and x can be expressed as

\[ \Delta y_t = \beta \Delta x_t - \lambda (y_{t-1} - y_0 - \gamma_1 x_{t-1}) + \nu_t \quad \text{and} \quad \nu_t \sim \text{NID}\left(0, \sigma^2\right) \]  

where \( \nu \) represents the normally and independently distributed (NID) random disturbances, also known as ‘white noise’ with the characteristics of mean zero, constant variance and zero autocovariances. The parameter \( \beta \) measures the short run effect on \( y \) of changes in \( x \), \( \gamma_1 \) measures the long run equilibrium relationship between \( y \) and \( x \),

\[ y_t = \gamma_0 + \gamma_1 x_t + u_t \quad \text{and} \quad u_t \sim \text{NID}\left(0, \sigma^2\right) \]  

\((y_{t-1} - y_0 - \gamma_1 x_{t-1})\) in equation (1) is the error correction term corresponding to lag version of (2), which represents the divergences from the long-run equilibrium. The parameter \( \lambda \) measures the extent of correction of such errors by adjustments in \( y \), the negative sign shows that adjustments are made towards restoring the long-run relationship. Short-run adjustments are therefore guided by, and consistent with, the long run equilibrium relationship.

2.2 Estimation procedure

2.2.1 Testing for order of integration

Before going on to the estimation of the error correction model, it is necessary to determine the order of integration of the variables considered for the model. Only variables of the same order of integration can be cointegrated and the existence of cointegration implies that there is a valid specification of the error correction model. A series \( Y_t \) is said to be integrated of order \( d \) if the series becomes stationary after differencing \( d \) times, denoted \( Y_t \sim I(d) \). For instance, if \( Y_t \) is stationary after differencing once, that is \( Y_t - \gamma_0 - \gamma_1 x_t \) is stationary, this is denoted as \( Y_t \sim I(1) \) and \( \Delta Y_t \sim I(0) \). The order of integration is commonly established using the Augmented Dickey Fuller test, where it is assumed that in each case the data generating process for each series is adequately represented by one of the three test equations below.

\[ \Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \alpha_2 t + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j} + \varepsilon_t \quad \text{and} \quad \varepsilon_t \sim \text{WN} \]  

\[ \Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j} + \varepsilon_t \quad \text{and} \quad \varepsilon_t \sim \text{WN} \]  

\[ \Delta Y_t = \alpha_1 Y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j} + \varepsilon_t \quad \text{and} \quad \varepsilon_t \sim \text{WN} \]  

where \( t \) is a deterministic time trend, \( p \) denotes the number of lags, and \( \varepsilon_t \) is an assumed Gaussian error term. The number of lagged terms (ie, \( p \)) is usually determined on the basis of Akaike (AIC) and Schwarz (SC) information criteria. The addition of \( p \) lagged terms in the
ADF equation is only to satisfy the Gaussian assumption for residual in (3). The null of non-stationary Y, series, i.e., $\alpha_1 = 0$, is tested against the alternative hypothesis of stationarity, $\alpha_1 < 0$. Equation (3) is estimated by using OLS results to calculate the test statistics for $\alpha_1$. Under the null $\alpha_1 = 0$, the test statistic is distributed not about zero but about a value less than zero. This reflects the fact that the OLS estimator $\hat{\alpha}_1$ is biased downwards. In this case, the Fuller (1976) Table 8.5.2 (p.373) is used for the asymptotic distribution of t-statistic (normally referred to as $\tau$) to compare the estimated t-statistic for $\alpha_1$. Using this $\tau$ value, it is possible to test the null hypothesis of non-stationarity despite the bias in the OLS estimator. Similarly, Dickey and Fuller (1981) Table III p.1062 is used to test $\alpha_2 = 0$, which represents the coefficient for the deterministic time trend in the ADF equation (3.1).

Nelson and Plosser (1982) provide evidence that, in fact, most economic time series do not display both deterministic and stochastic trends; that is, a significant time trend is most unlikely when unit root is present. Lloyd and Rayner (1993), however, argued that detecting the correct form of non-stationary behaviour is important since a trend stationary process can not be co-integrated with a difference stationary process, whereas a stochastic trend is difference stationary and a deterministic trend is a trend stationary process. In the ADF equation (3.1), the joint null such as $\alpha_1 = \alpha_2 = 0$ can be tested by the F-statistic. However, since under the null hypothesis of a stochastic trend conventional distribution theory does not hold, critical values for the test statistic are obtained from Dickey and Fuller (1981) Table 6, p. 1063 (the F like test in this case is generally refereed as $\phi_3$ statistic). Failure to reject the above joint null hypothesis implies Y is subject to a stochastic trend.

However, all these critical values which we used to test the hypothesis in ADF are only valid if $Y_t$ is a purely autoregressive process. If a moving average process is present in $Y_t$, the critical values of the ADF test presented above are inappropriate and an alternative test should be used (see, Leybourne, 1992). In addition, if the residuals from the ADF test equation are not white noise, an assumption of the test is violated. In this case, visual inspection of a correlogram of the ADF test equation residuals can be helpful. Hence, the low power of the ADF test for unit roots should be considered with caution and should not be regarded as precise. The critical values used in this test are only approximate guides. Therefore, a sufficiently higher test statistic than the critical value should be obtained to ascertain the validity of the test results. Alternatively, the Phillips-Peron (1988) unit root test can be applied because it uses a non-parametric correction for serial correlation and may improve the power of the test.

2.2.2 Testing for co-integration

The validity of an error correction model (ECM) rests on the concept that there exists a long run equilibrium relationship between the relevant series. Co-integration tests answer whether or not there exists such a relationship. Before using the co-integration test, it is necessary to establish whether each series is integrated to the same order. This can be done by using the ADF test for unit roots as described above. If it is found that all the series of interest are I(1)-that is they become stationary on first differencing, then the co-integration test is completed by using the following Augmented Dickey Fuller test.
First, the hypothesised equilibrium relationship, say between $X$ and $y$ in (2), is estimated by using OLS (which is known as the static co-integration regression). The residuals $\hat{u}_t$ from this regression are obtained as,

$$\hat{u}_t = Y_t - \gamma_0 - \gamma_1 x_t$$

(4)

The test for no co-integration is then given by applying the following augmented Dickey-Fuller regression.

$$\Delta \hat{u}_t = \alpha \hat{u}_{t-1} + \sum_{j=1}^{p} \phi_j \Delta \hat{u}_{t-j} + v_t$$

(5)

where $v_t$ is an assumed Gaussian error. The null $H_0: \alpha = 0$ against $H_a: \alpha < 0$ is tested by using ADF critical values for co-integration provided in Engle and Yoo (1987), Table 3. More importantly, equation (5) does not include an intercept or a time trend, since the $u_t$ must have a zero mean and it is not expected that they would exhibit a deterministic trend.

The above test would be a valid approach if the true parameter values in (2), and hence $u_t$ are known. Unfortunately, only OLS estimates are available for $\gamma_0, \gamma_1$ and $u_t$. As Engle and Granger (1987) pointed out, OLS minimises the residual sum of squares, and gives parameter estimates that are most likely to result in stationary residuals. The normal Dickey-Fuller procedure will therefore reject the null of a unit root more often than it should. Therefore, the t-ratio for the OLS estimate of $\alpha$ in (5) needs to be more negative than suggested by the ADF critical before safely rejecting the null of a unit root. Engle and Granger consider seven possible statistics for testing stationarity for $u_t$ from (2). They eventually recommend the ADF test based on (5) but with different critical value as mentioned above.

Alternatively, Engle and Granger suggested the following test for co-integration.

$$\hat{u}_t = \phi \hat{u}_{t-1} + v_t$$

(6)

In this test, $H_0: \phi = 1$ is tested against $H_a: \phi < 1$. In other words, the null of no co-integration is tested. But (6) implies that residual $u_t$ from the co-integration-regression (2) follows a first order autoregressive process. This problem can be easily tackled by using Durbin Watson test since when $\phi = 1$, the DW statistic approaches zero, otherwise it is greater than zero. The critical values of the test statistic are given in Engle and Yoo (1987). A large DW test statistic is suggestive of I(0) residuals and evidence of a co-integrating relationship. This test for stationarity is also known by co-integrating regression Durbin-Watson (CRDW) test. The CRDW as proposed by Bhargava (1983) should only be used for obtaining a “quick approximate result” (Engle and Granger, 1987, p. 269). The CRDW test works well when the disturbances in the co-integration regression follow a first-order autoregressive process, otherwise, it has very different critical values for alternative specifications. Hence, the ADF test for co-integration is superior compared to the CRDW test since the former accommodates higher order autoregressive (AR) processes. However, the ADF result should be judged and interpreted carefully due to its low power against near-unit-root processes.
2.2.3 Estimation of the Error Correction Model (ECM)

As mentioned in the outset, an ECM implies the existence of an underlying equilibrium relationship in the long run. Engel and Granger (1987) prove that if two variables are co-integrated (that is, if an equilibrium relationship exists) then the short-run 'disequilibrium' relationship between the two variables can always be represented by an ECM. Indeed, if there is no equilibrium relationship, then the short run behaviour can not be represented by ECM. This ultimately requires that the first step in ECM estimation is to test for co-integration as mentioned in section 2.2.2. If \( x \) and \( y \) are co-integrated, then an ECM \((1)\) can be estimated by using either of the procedures mentioned below.

Engle and Granger propose a two-step procedure for the estimation of an ECM. This starts with OLS estimation of a co-integrating regression such as \( (2) \) and then testing for co-integration. If the null hypothesis of no co-integration is rejected, the second step uses the lagged residuals from the co-integrating regression as the error correction term in an error correction model such as \( (1) \), thus imposing the long-run relationship as a restriction. Thus, the ordinary least squares technique is employed to estimate a regression such as \( (7) \) in the second stage.

\[
\Delta y_t = \beta \Delta x_t - \lambda \hat{u}_{t-1} + v_t \quad \text{and} \quad v_t \sim 'WN' \Rightarrow \text{NID} \left(0, \sigma^2 \right) \tag{7}
\]

Although the OLS estimators from the co-integrating regression possess the asymptotic properties of consistency and are highly efficient, the small sample biases in the parameter estimates for the co-integrating regression may be substantial (Stock, 1987; Banarjee et al., 1986), and these may carry over to the estimate of the error correction parameter \( (\lambda) \) in the second stage. This bias appears to be inversely related to the goodness of fit, so a high \( R^2 \) at this stage may be necessary to obtain acceptable results from the two-step procedure (Hallam and Zanoli, 1992).

An alternative to the Engle and Granger two-step procedure is to apply OLS to \( (8) \) directly and hence estimate both short-run and long-run parameters together (Wickens and Breusch, 1988).

\[
\Delta y_t = \delta + \beta \Delta x_t - \lambda_1 \hat{y}_{t-1} + \lambda_2 \gamma_1 x_{t-1} + v_t \quad \text{and} \quad v_t \sim 'WN' \Rightarrow \text{NID} \left(0, \sigma^2 \right) \tag{8}
\]

where \( v_t \) is the assumed Gaussian error and \( \delta = \lambda \star \gamma_0 \). The estimate of the long-run parameter \( \gamma_2 \) can then be calculated as the ratio of coefficients of \( x_{t-1} \) and \( y_{t-1} \). Likewise an estimate of \( \gamma_0 \) is obtained from the ratio of the constant term, \( \delta \) to the coefficient of \( y_{t-1} \). This approach is also not without criticism. It is argued that the short-run parameter is treated equivalently to the long-run parameter in estimating parameters in \( (8) \), which actually should not be the case. The long run parameters are often driven by theory while the short-run parameters may not be related to the theory. However, there is some evidence that the small sample bias is smaller for the latter estimator than it is with the Engle-Granger two-step procedure.
As an alternative to both of the methods mentioned above, the ECM used in this paper is estimated using Hendry's general to specific approach, which is based on an Autoregressive Distributed Lag (ADL) model (Hendry, 1995) as follows

$$y_t = \mu + \sum_{i=1}^{k_1} \delta_{1i} y_{t-i} + \sum_{i=0}^{k_2} \delta_{2i} x_{1,t-i} + \sum_{i=0}^{k_3} \delta_{3i} x_{2,t-i} + \cdots + \sum_{i=0}^{k_4} \delta_{3i} x_{3,t-i} + \cdots + u_t \quad \cdots \quad (9)$$

where $u_t$ is a zero-mean, white noise error term and $\delta_{ji}$ depicts the short run dynamic adjustment parameter for the jth variable. The implied long run equilibrium solution to (9) can be expressed as

$$y_t^* = \phi_1 x_{1t} + \phi_2 x_{2t} + \phi_3 x_{3t} + \cdots \quad .$$

The long run equilibrium parameters are then given by

$$\phi_1 = \frac{\mu}{1 - \sum_{i=1}^{k_1} \delta_{1i}}; \quad \phi_2 = \frac{\sum_{i=0}^{k_2} \delta_{2i}}{1 - \sum_{i=1}^{k_1} \delta_{1i}}; \quad \phi_3 = \frac{\sum_{i=0}^{k_3} \delta_{3i}}{1 - \sum_{i=1}^{k_2} \delta_{2i}}; \quad \phi_4 = \frac{\sum_{i=0}^{k_4} \delta_{3i}}{1 - \sum_{i=1}^{k_3} \delta_{3i}}; \quad \cdots$$

The 'residual' term $ZD_t = (y_t - y_t^*)$ is the computed error correction series. The stationarity of $ZD_t$ then implies that the four series under consideration are linked in a long run cointegrating relationship.

3. Variable selection and data set

The determinants of land price are still a subject of discussion (see for example, Lloyd and Rayner 1990; Hallam et al., 1992). However, there seems to be an emerging consensus on the importance of returns to farmland in determining land prices (for example, Melichar, 1979; Alston, 1986; Burt, 1986; Lloyd and Rayner, 1990), which stems from the Ricardian rent theory. But on prior grounds other variables can be eliminated as potential contributing factors to past growth in land prices. The work that follows is an attempt to establish the empirical importance of the competing explanations.

The most conventional hypothesis is that growth in land prices arises from growth of income to land. There are differences of opinion in the selection of the most appropriate proxy for farm returns. Net farm income was used in the original Traill (1979) model, but it has been suggested that this is not an appropriate proxy of the return to land alone (Melichar, 1979). There is considerable support of rent as an appropriate proxy for farm returns (for example: Alston, 1986; Burt, 1986; Lloyd and Rayner, 1990). However, it has been argued that the use of rents is of limited relevance in the New Zealand case where most land is owner occupied and owner operated. The very limited market for rental land can not reflect the measure of returns to farmland. In their New Zealand study, Seed and Sandrey (1985) used residual
income as the proxy for return to farmland. They obtained the residual income by adding depreciation to net farm income and deducting managerial reward. They have not added back interest payments to the net farm income. Melichar (1979) argued that interest paid on farm debt should also be included in the net farm income, as the goal is to determine the returns to assets rather than to equity. Therefore, this study uses the net residual income as the proxy for farm return, which is derived by adding depreciation and interest to net income and deducting managerial salary. This paper adopted the definition of net farm income as defined by the Meat and Wool Board’s Economic Service in its Sheep and Beef Farm Survey publications. Accordingly, net farm income is the difference between the gross farm income and gross expenditure including managerial salary, depreciation, and interest paid to debt but not taxes. Since sheep, beef, and dairy farms covered more than 85% of farmland in New Zealand agriculture, the average farm income obtained by accounting for these sectors is assumed to provide proxy average return for all farmland. In this paper, the proxy for average annual net income for farmland is calculated as the mean of net residual income for all class sheep and beef farms, and the annual net residual income for owner operated factory supply dairy farms. The average net residual income series is calculated for a hectare of effective farmland area. All time series data sets are obtained from the New Zealand Meat and Wool Board’s Economic Service, Sheep and Beef Farm Survey, and the Dairy Board’s Economic Survey of Factory Supply Dairy Farms.

Another less conventional hypothesis that has appeared frequently in the economic literature is that inflation causes real growth in land prices. Feldstein (1980) argued that sustained growth of land prices could arise from a continual increase in the inflation rate. An upward trend in the rate of inflation might add to the effect of growth of income to land in causing growth in land prices. In his U.S. case study, he reported that increases in the anticipated inflation rate cause increases in real land prices because of the characteristics of the U.S. tax system. Alston (1986) has found empirical evidence against Feldstein’s hypothesis. On the contrary, he found the empirical evidence in favour of the more conventional hypothesis originally proposed by Melichar, that is the growth in land price arises from growth in net rental income to farmland. He further developed the counter hypothesis that increases in expected inflation have a negative effect on real land prices, though the effect of inflation has been comparatively small. Burt (1986) also found no effect of inflation on movement of real land prices. However, Lloyd and Rayner (1990) have found an effect of general inflation on real land price and developed a modified version of a model that includes the inflation as the explanatory variable to determine the land price in UK case. Here, the effect of changes in general price inflation on land value is examined particularly in the New Zealand context. This study assumes the GDP deflator index as the proxy for inflation, data for which are obtained from the IMF publication ‘International Financial Year Book’.

Moreover, the rationale of considering land as an investment asset leads to the inclusion of interest rates as a likely variable to explain the movement of land prices. Further support for the inclusion of the real interest rate is provided by Burt (1986). According to Burt, ‘With the long term investment characteristics of farm land and the sizeable transaction costs involved, market participants are apt to use an estimated long run equilibrium rate of interest in the classic capitalisation formula to approximate land values.’ Lloyd, Rayner, and Orme (1991) argued that real interest rates as an opportunity cost to capital investment may have short run influence on market prices. This paper examines the possible role of the real interest rate on the movement of farmland prices in New Zealand. The return on short term government
bonds is considered as the proxy of the interest rate, and is obtained from the selected issues of the Reserve Bank Bulletin, Wellington.

The inclusion of other variables such as taxes was not considered in this study. As Burt (1986) notes, 'Taxes have an obvious influence on land prices.' But '...taxes or tax rates are not necessarily required as independent variables in a time series regression equation for land prices' (Burt, 1986). Lloyd et al. (1991) further notes that 'taxes and tax rates probably have only second order effects on real land prices'.

Time series data for average farmland prices came from Valuation New Zealand. This represents the average per hectare farm land prices based on the open market sales of freehold rural farmland. The time series for all variables were taken for the period 1970 to 1997. All economic variables of interest used to explain land price appreciation were adjusted for general price inflation before analysis using the GDP deflator index.

Most empirical time series exhibit a variation that increases in both the mean and dispersion in proportion to the absolute level of the series. For example, the real land price series (depicted in Figure 1.b) evolved through time, revealed that both the mean and variance increased. Though the application of the difference operator frequently removes a time-dependent mean, it has a negligible effect on stabilising the variance of empirical time series. This ultimately motivates the use of power transformation in empirical work; the logarithmic transformation is generally adequate in practice, although the type of transformation required will depend ultimately on the severity of the trend in the variance. As a rule of thumb, any I(1) series should be transformed into logarithms although it is advisable whether to test for the appropriate transformation prior to econometric analysis (see Mills, 1990, pp.48-50). All series used in this study were transformed to natural logarithmic form before conducting any tests.

4. An error correction model for farmland price determination

In this section, the error correction model for farmland price determination is developed, estimated, and tested by following the Hendry (1995) two-step procedure as outlined above. The basis of the analysis is the following dynamic error correction model

\[
\Delta LLP_t = \beta_3 \Delta LNRRI_t + \beta_2 \Delta LRIR_t + \beta_1 \Delta LINF_t
\]
\[+ \lambda (LLP_{t-1} - \gamma_0 - \gamma_1 LNRRI_{t-1} - \gamma_2 LRIR_{t-1} - \gamma_3 LINF_{t-1}) + \nu_t \]

(10)

where \( v_t \) is a Gaussian error term,
LLP = log of real land price ($ per hectare)
LNRRI = log of real net residual income in $
LRIR = log of real interest rate
LINF = log of inflation rate, and \( \Delta \) denotes the first difference operator.

The validity of the error correction formulation in (10) requires the existence of a long run relationship or co-integration amongst the series concerned. However, co-integration requires that the series concerned are integrated of the same order and that a linear combination of
these variables, described by a cointegrating regression, is integrated of an order less than each individual series. The modelling strategy therefore starts with tests for the order of integration using the Augmented Dickey Fuller test as described in section 2.2.1, and then tests for the existence of a co-integrating vector involving the four variables of interest: real farmland prices, real net farm residual income, real interest rate, and inflation. For the integration tests, it is assumed that the data generating process for each of the four series is adequately represented by one of the three forms (3.1, 3.2, or 3.3) presented on page 6 above.

The testing procedure begins by estimating equation (3.1) for each series using ordinary least squares. Estimating equation (3.1) for an individual original series has relative advantages over (3.2) and (3.3) in that the null hypothesis of unit root as well as the stochastic trend can be tested simultaneously. The hypothesis that the series concerned has a unit root with stochastic trend \((\alpha_1=0\) and \(\alpha_1=\alpha_2=0\)) is tested using the Dickey Fuller \(\tau\) and \(\phi_3\) statistic. The estimated and critical values of \(\tau\) and \(\phi_3\) for all variables are presented in Table 1a.

### Table 1a

<table>
<thead>
<tr>
<th>Series</th>
<th>Ho: (\alpha_1=0)</th>
<th>Ho: (\alpha_1=\alpha_2=0)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>t-statistic</td>
<td>(\tau^c) at 5%</td>
</tr>
<tr>
<td>LLP(_t)</td>
<td>-1.7448</td>
<td>-3.41</td>
</tr>
<tr>
<td>LNRI(_t)</td>
<td>-1.2137</td>
<td>-3.41</td>
</tr>
<tr>
<td>LIRIR(_t)</td>
<td>-1.1267</td>
<td>-3.41</td>
</tr>
<tr>
<td>LINF(_t)</td>
<td>-1.1391</td>
<td>-3.41</td>
</tr>
</tbody>
</table>

The ADF test result suggests that the null of a unit root in each series can not be rejected at 5% significance level since the critical value of the \(\tau\) statistic is more negative than the calculated \(t\) value. The test results for joint null \(\alpha_1=\alpha_2=0\) also provides evidence in its favour. Though the use of the Dickey-Fuller critical values to compare the test statistics has been criticised on its empirical validity (see Leybourne, 1992), all series have passed the purely autoregressive process before carrying out the ADF test. Therefore, the test results presented above provide confirmatory evidence that all individual series follow a stochastic trend instead of a deterministic trend. As a rule, all time series with a stochastic trend may be cointegrated. To have confirmatory evidence that all individual variables are I(1), the same ADF test is carried out on their first differences because co-integration requires the same order of integration for each series. The test results are presented in Table 1b.
Table 1b

Unit Root Test Results (differences)

<table>
<thead>
<tr>
<th>Series</th>
<th>Ho: $\alpha_t=0$</th>
<th>$t$-statistic</th>
<th>$\tau^2$ at 5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta \ln \text{INF}_t$</td>
<td>-2.7114</td>
<td>-3.41</td>
<td></td>
</tr>
<tr>
<td>$\Delta \ln \text{LP}_t$</td>
<td>-2.6641</td>
<td>-2.57</td>
<td></td>
</tr>
<tr>
<td>$\Delta \ln \text{NRI}_t$</td>
<td>-3.149</td>
<td>-2.57</td>
<td></td>
</tr>
<tr>
<td>$\Delta \ln \text{RIR}_t$</td>
<td>-2.8297</td>
<td>-2.57</td>
<td></td>
</tr>
<tr>
<td>$\Delta^2 \ln \text{INF}_t$</td>
<td>-4.0627</td>
<td>-2.57</td>
<td></td>
</tr>
</tbody>
</table>

The above ADF test results were obtained from the regression equation of (3.2) for each individual series. The results revealed that all series in the first difference but inflation in the second difference are stationary at the 5% significance level. The non-rejection of joint null $\alpha_1=\alpha_2=0$ for $\Delta \ln \text{INF}_t$ provides confirmatory evidence that the series shows a stochastic trend instead of a deterministic trend. Hence, each series: LLP, LNRI, LRIR, and $\Delta \ln \text{INF}_t$ is considered to be I(1), thus they may be co-integrated.

To test for cointegration we firstly estimate an appropriate autoregressive distributed lag (ADL) model, compute the implied long run equilibrium coefficients and the associated error correction term (ZD), then test ZD for a unit root. The AIC criterion selected an ADL(2,2,1,0) specification, the OLS results for which (obtained from Microfit 4.0) are presented below.

Table 2

(a) Autoregressive Distributed Lag Estimates
ADL(2,2,1,0) selected based on Akaike Information Criterion
Dependent variable is LLP. 25 observations used for estimation from 1973 to 1997

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>$T$-Ratio[p-value]</th>
</tr>
</thead>
<tbody>
<tr>
<td>LLP$_{t-1}$</td>
<td>0.78980</td>
<td>0.1495</td>
<td>5.2831[0.000]</td>
</tr>
<tr>
<td>LLP$_{t-2}$</td>
<td>-0.28403</td>
<td>0.1492</td>
<td>-1.9033[0.075]</td>
</tr>
<tr>
<td>LNRI$_t$</td>
<td>0.19720</td>
<td>0.1497</td>
<td>1.3170[0.206]</td>
</tr>
<tr>
<td>LNRI$_{t-1}$</td>
<td>-0.12081</td>
<td>0.1668</td>
<td>-0.72424[0.479]</td>
</tr>
<tr>
<td>LNRI$_{t-2}$</td>
<td>0.31631</td>
<td>0.1500</td>
<td>2.1005[0.052]</td>
</tr>
<tr>
<td>LRIR$_t$</td>
<td>0.68168</td>
<td>0.15731</td>
<td>4.3335[0.001]</td>
</tr>
<tr>
<td>LRIR$_{t-1}$</td>
<td>-0.64413</td>
<td>0.12504</td>
<td>-5.1514[0.000]</td>
</tr>
<tr>
<td>$\Delta \ln \text{INF}_t$</td>
<td>0.17971</td>
<td>1.6803</td>
<td>0.10695[0.916]</td>
</tr>
<tr>
<td>Const</td>
<td>1.5040</td>
<td>0.6212</td>
<td>2.4212[0.028]</td>
</tr>
</tbody>
</table>

R-Squared | 0.97587 | R-Bar-Squared | 0.96381 |
S.E. of Regression | 0.16668 | F-stat(8,16) | 80.8862[0.000] |
Residual Sum of Squares | 0.44449 | Equation Log-likelihood | 14.8977 |
Akaike Info. Criterion | 0.58977 | DW-statistic | 0.15000 |
(b) Diagnostic Tests for ADL(2,2,1,0) Model

<table>
<thead>
<tr>
<th>Test Statistics</th>
<th>LM Version</th>
<th>F Version</th>
</tr>
</thead>
<tbody>
<tr>
<td>A: Serial Correlation</td>
<td>CHSQ(1) = 2.0149[0.156]</td>
<td>F(1,15) = 1.3149[0.269]</td>
</tr>
<tr>
<td>B: Functional Form</td>
<td>CHSQ(1) = 0.15719[0.692]</td>
<td>F(1,15) = 0.09491[0.762]</td>
</tr>
<tr>
<td>C: Normality</td>
<td>CHSQ(2) = 2.5568[0.278]</td>
<td>Not applicable</td>
</tr>
<tr>
<td>D: Heteroscedasticity</td>
<td>CHSQ(1) = 2.2648[0.132]</td>
<td>F(1,23) = 2.2912[0.144]</td>
</tr>
</tbody>
</table>

*Note: figures in parentheses are p-values for the corresponding test statistics.*

The individual coefficient values shown in the first part of the table are not of particular interest due to likely multicollinearity amongst the lagged variables, but the diagnostic tests in the second part do not give any indication that the ADL(2,2,1,0) model is misspecified. Hence these results are used to solve for the long run equilibrium coefficients \( \hat{\theta}_j \) and their asymptotic variances as discussed on page 8. The results are displayed below.

Table 3

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>T-Ratio[p-value]</th>
</tr>
</thead>
<tbody>
<tr>
<td>LNRII</td>
<td>0.79457</td>
<td>0.29969</td>
<td>2.6513[0.017]</td>
</tr>
<tr>
<td>LRIR</td>
<td>0.075968</td>
<td>0.24886</td>
<td>0.30527[0.764]</td>
</tr>
<tr>
<td>ΔLINF</td>
<td>0.36361</td>
<td>3.3947</td>
<td>0.10711[0.916]</td>
</tr>
<tr>
<td>Const</td>
<td>3.0432</td>
<td>1.0700</td>
<td>2.8441[0.012]</td>
</tr>
</tbody>
</table>

For the three proposed explanatory variables, notice that only LNRII has a statistically significant long run coefficient. The value of this elasticity coefficient indicates that, in the long run, a 1% increase in real net farm residual income will result in an increase of about 0.8% in real farm land prices. The estimated long run equilibrium values for the log of real land prices associated with these long run coefficients are calculated from the following relation

\[
LLPF_t = 3.0432 + 0.79457*LNRII_t + 0.075968*LRIR_t + 0.36361*ΔLINF_t
\]

As shown in Figure 2, there is a close relationship between the actual values of the real land price (LLP) and the fitted values (LLPF). In fact, the coefficient of linear correlation between LLP and LLPF is 0.955. In each period the algebraic difference between the value of LLP and LLPF measures the extent to which LLP deviates from its estimated long run equilibrium value for that period. That is, whenever LLP is greater (less) than LLPF, actual farm land prices are above (below) their long run equilibrium values.
The computed values of the error correction term \((ZD_t = LLP_t - LLP_F_t)\) provide a measure of long run disequilibrium in the real price level for farm land. Notice that the time plot of \(ZD_t\), shown below does not indicate that this series is nonstationary. That is, the plot in Figure 3 suggests that the four variables of interest may be linked in a long run cointegrating relationship with parameter values given in Table 3 above.
Turning now to the issue of cointegration, the Representation Theorem of Engle and Granger (1987) states that the existence of a valid error correction model (ECM) implies cointegration. An ECM representation of the adjustment process for the 4 series of interest is:

\[ \Delta \text{LLP}_t = \beta_1 \Delta \text{LNRI}_t + \beta_2 \Delta \text{LRIR}_t + \beta_3 \Delta^2 \text{LINF}_t + \lambda \text{ZD}_{t-1} + \nu_t \quad \text{where } \nu_t \sim 'WN' \quad \text{(12)} \]

which will qualify as a valid ECM if it is the case that \(-1 < \lambda < 0\), provided the residual term has a Gaussian distribution. Hence equation (12) is estimated with OLS and the t-statistic for \(\hat{\lambda}\) is used to test the null hypothesis of no cointegration. Table 4 shows the results obtained.

### Table 4

(a) Ordinary Least Squares Estimation of the Error Correction Model
Dependent variable is \(\text{LLP}\). 25 observations used for estimation from 1973 to 1997

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>T-Ratio [p-value]</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\Delta \text{LNRI}_1)</td>
<td>0.30437</td>
<td>0.12101</td>
<td>2.5153 [0.020]</td>
</tr>
<tr>
<td>(\Delta \text{LRIR})</td>
<td>0.63452</td>
<td>0.11741</td>
<td>5.4043 [0.000]</td>
</tr>
<tr>
<td>(\Delta^2 \text{LINF})</td>
<td>-0.40870</td>
<td>1.1357</td>
<td>-0.3599 [0.723]</td>
</tr>
<tr>
<td>(\text{ZD}(-1))</td>
<td>-0.41745</td>
<td>0.11352</td>
<td>-3.6772 [0.001]</td>
</tr>
</tbody>
</table>

R-Squared 0.96857 R-Bar-Squared 0.96408
S.E. of Regression 0.16963
F-stat(3,21) 215.7159 [0.000]
Mean of Dependent Variable -0.0076362 S.D. of Dependent Variable 0.89501
Residual Sum of Squares 60425 Equation Log-likelihood 11.0595
DW-statistic 1.4856

(b) Diagnostic Tests

<table>
<thead>
<tr>
<th>Test Statistics</th>
<th>LM Version</th>
<th>F Version</th>
</tr>
</thead>
<tbody>
<tr>
<td>A: Serial Correlation</td>
<td>CHSQ(1)= 2.4024[0.121]</td>
<td>F(1,20)= 2.1263[0.160]</td>
</tr>
<tr>
<td>B: Functional Form</td>
<td>CHSQ(1)= 0.2605[0.610]</td>
<td>F(1,20)= 0.2106[0.651]</td>
</tr>
<tr>
<td>C: Normality</td>
<td>CHSQ(2)= 0.4626[0.793]</td>
<td>Not applicable</td>
</tr>
<tr>
<td>D: Heteroscedasticity</td>
<td>CHSQ(1)= 0.7941[0.373]</td>
<td>F(1,23)= 0.7545[0.394]</td>
</tr>
</tbody>
</table>

A: Lagrange multiplier test of residual serial correlation
B: Ramsey's RESET test using the square of the fitted values
C: Based on a test of skewness and kurtosis of residuals
D: Based on the regression of squared residuals on squared fitted values
As none of the diagnostic tests indicates statistical inadequacy, we may test the validity of the error correction model by testing the null hypothesis $H_0: \lambda = 0$, the rejection of which indicates that equation (12) can be interpreted as a valid error correction model. From Table 4 we see that $\hat{\lambda}$ is negatively signed and the ratio of $\hat{\lambda}$ to its standard error (-3.6772) is large in absolute value. However, this ratio does not have the standard t-distribution but Banerjee, Dolado and Mestre (1993, Table 4) have tabulated critical values for a variety of specifications of the cointegrating vector. For a vector consisting of 4 series plus a constant, and with a sample of 25 observations, the 10% critical value is -3.68.

These results, together with the evidence provided in Figure 3, are sufficient to indicate that the four series of concern are linked in a cointegrating relationship. Hence, we may make the following interpretations of the estimated coefficients.

For the short run adjustment process, the estimated coefficients for the change in net residual income ($\Delta LRI_{11}$) and the change in the real interest rate ($\Delta LRI_R$) are statistically significant at conventional levels of significance, but the coefficient for the change in the inflation rate ($\Delta^2 LINF$) is not. This implies that real farm income and the real interest rate both have significant and positive impacts on the change in real farm land prices whereas inflation has a negative but very negligible effect. As the dependent variable and all short-run variables are presented in the first difference of the natural logs, each coefficient value also relates the growth rate of real land prices to the percentage changes in the corresponding explanatory variable (see Lloyd and Rayner, 1990, 1993 for more explanation). For instance, $\beta_1=0.30437$ implies that in the short run, if the growth in farm real income increases up by 1% at time $t$, the growth in land price increases by 0.30% in the same period. Likewise, if the growth in real interest changes by 1 percentage point, this produces a 0.63% change in the growth of real land value in the same direction. If changes occurred in both variables by 1% in each, the growth in real land value changes by 0.93%, keeping effects of other variables constant over the study period. Moreover, in the long run the real land value will correct its new equilibrium value by almost 42%. In another words, when actual land price deviates from its estimated long run equilibrium value, 42% of this discrepancy or equilibrium error will be corrected in the following next year. This behaviour might be interpreted as a tendency for the market to recognise its mistakes and a corrective adjustment is made the following year.
5. Conclusion
The principal aim of this paper has been to test for both short-run and long-run effects of microeconomic, macroeconomic and institutional factors related to the historical movement of farmland prices in New Zealand. It has used the error correction model (ECM) corresponding to the statistical concept of co-integration. It has been shown here that only real net farm residual income has a long run effect on real land price movements, which ultimately supports the hypothesis originally proposed by Melichar (1979). Neither inflation nor the real interest rate has any significant effect on it. The empirical evidence, however, indicates that the real interest rate has a significant effect on the movement of real land price in the short-run, which is consistent with the hypothesis of Lloyd, Rayner and Orme (1991). The finding of a negative but statistically insignificant short-run effect of inflation on farmland price movements provides evidence against Feldstein (1980) but in support of Alston (1986). The practical implication of this finding can be interpreted as follows: investors in the immediate short-run may be more willing to invest in farmland when the real interest rate is very low thus explaining nearly two thirds of the reduction in land price. This may stimulate demand in farmland market. On the contrary, higher real interest rates and higher real farmland values may cause low demand pressure in the farmland market. The capitalised value of farmland investment is evaluated as one third of the annual rate of residual income in the immediate short-run. However, the long-run equilibrium value of real farmland price is solely determined by the net residual income on farmland. In other words, the annual capitalised value on farmland investment is approximately equivalent to the annual rate of residual income on farmland investment in the long-run. Investors will not discount the land value to compensate for inflationary pressures. This may be due to the average low inflation rate over the study period.

6. Limitations
Many problems are exacerbated in using aggregate data to analyse determinants of land prices. Such problems may arise due to extreme heterogeneity in land quality, and exclusion of nonagricultural values of farmland in historical farm income. Further, inaccurate estimates of farm income might have resulted from accounting and sampling errors in the historical observations.
References:


International Monetary Fund. *International Financial Statistical Year Book.* (selected issues only).


New Zealand Meat and Wool Board’s Economic Service. *The New Zealand Sheep and Beef Farm Survey.* (selected issues only).


