University of Wisconsin-Madison
Department of Agricultural & Applied Economics

December 2002 **Staff Paper No. 457**

On the Decline of Agriculture in Developing Countries: A Reinterpretation of the Evidence

By

Chaiyuth Punyasavatsut and Ian Coxhead

AGRICULTURAL & APPLIED ECONOMICS

STAFF PAPER SERIES

On The Decline of Agriculture in Developing Countries: A Reinterpretation of the Evidence

Chaiyuth Punyasavatsut
Thammasat University

Ian Coxhead*
University of Wisconsin-Madison

Abstract

Conventional explanations for the relative decline of agriculture in developing countries stress secular, demand-side phenomena, specifically Engel effects. This view has been challenged by quantitative analyses emphasizing supply-side effects such as differences in factor endowment growth rates. The innovation in this paper is to investigate the extent to which agricultural decline is in fact generated by policies rather than by fundamental preference or endowment shifts. Econometric results using Thai data indicate that policies are strongly influential, but that the direction and strength of influence varies over time. We explore implications for the interpretation of past development strategies and future policy formation.

JEL classifications: O11, O13, Q18, O53.

Key words: Growth, structural change, agricultural policy, Thailand

---

* Corresponding author. Address for correspondence: 412 Taylor Hall, University of Wisconsin-Madison, 427 Lorch St., Madison, WI 53706. Phone: 608/262-6390. Fax: 608/262-4376. Email: coxhead@facstaff.wisc.edu

1. Introduction

The decline of agriculture's contribution to national income is a central feature of economic development. The share of GDP originating in agriculture, initially greater than 50% in many poor countries, is typically much less than 10% in wealthy countries. Time-series data from individual developing countries exhibit the same trend. In the course of this shift, agriculture releases primary factors to sectors making more intensive use of reproducible capital, thus contributing to faster economic growth. In order to understand growth and development processes, therefore, it is necessary to examine the forces that govern structural change. Analysis of the determinants of agriculture's decline should also inform the interpretation and evaluation of the effects of sector policies on growth in developing economies.

Given the importance of the subject, it is surprising to find that very little effort has been devoted to quantifying the sources of agricultural decline. To our knowledge the only empirical studies are the two by Martin and Warr (1993, 1994), which we discuss in detail below. This paper makes three additional contributions. First, we quantify the effects of major price policies as determinants of structural change separately from those of underlying price trends. Second, we identify separately the influence of human capital accumulation as a factor in structural change. Third, we present findings indicating that over time, the sources of agricultural decline may shift and even change sign as economies and policies evolve.

The paper is organized as follows. Section 2 discusses the theoretical background and new elements introduced in our analysis. Section 3 describes the basic model. Section 4 discusses estimation issues, Section 5 presents data and measurement, and Section 6 presents empirical results. A final section provides concluding observations on the development process and development policies.

2. Theory and empirical motivation

Sources of structural change

The "classics" of development economics ascribe the relative decline of agriculture to fundamentally secular forces, specifically asymmetric changes in preferences, technologies, and factor endowments (Lewis 1954; Johnston and Mellor 1961; Chenery *et al*. 1986; Timmer 1988).[1]

On the demand side, less than unitary elasticities of demand for staple foods mean that as per capita income grows, a declining proportion of household expenditure is devoted to agriculture. By Engel's law, as per-capita income rises, expenditure shifts toward services and manufactured goods relative to food. In a closed economy, the preference shift causes the relative price of food to decline, other things equal. This in turn reduces returns to factors used in agricultural production, causing a net migration of labor and capital to other sectors (Schultz 1953; Timmer 1988), thus reducing relative growth rates of agricultural output and employment. In a small open economy with both tradable and non-tradable sectors, agriculture's share in GDP also declines if demand for non-tradable goods is income-elastic, since demand growth bids up their prices relative to those of goods more freely traded in world markets (Anderson 1987).

On the supply side, explanations for agricultural decline have been dominated by asymmetric growth in factor endowments and differences in sectoral productivity growth rates. The two are closely related; for the Hicks-neutral case, differences in technical change rates between sectors affect the composition of aggregate output in a way exactly analogous to differential rates of factor endowment growth. Curiously, however, it is technical progress rather than factor endowment growth that has received more attention in the development economics literature.[2] The effects of unbalanced endowment growth rates on economic structure are nevertheless potentially profound. This prediction is familiar in the form of Rybczynski's theorem, which states that in a two-sector, two-good economy, an increase in the total supply of capital relative to labor will influence the output mix, for given product prices and technology. Specifically, the theorem predicts that a small increase in the capital-labor ratio will increase the output of the relatively capital-intensive good, and reduce the output of the relatively labor-intensive good.[3] Given that an increase in the capital-labor ratio, fueled by savings and investment, is the mainspring of structural change in all classical models of growth, the preference for technical progress stories over endowment growth in supply-side explanations of agricultural decline is puzzling.

It is only very recently that the foregoing assertions about the sources of structural change have been subjected to any formal quantitative testing. Recent econometric analyses have suggested that it is changes in factor endowments—the so-called Rybczinski effects—that actually dominate, with relative price changes (i.e. demand side effects) playing a much smaller role, and technical progress a negligible one. This ordering reverses that of earlier writing. As the authors of these analyses note, their results "appear to require a reorientation of the literature on the declining share of agriculture in open economies" (Martin and Warr 1993:398).

Whatever the true distribution of explanatory power among the three sources identified above, the overall emphasis on secular forces suggests a uniformity in the sources of agricultural decline that is not matched by developing country experience. Growth and structural change differences among countries, and over time within countries, are most frequently explained in terms of economic policies. These include agricultural policies, but also those directed at other areas of economic activity but which affect relative sectoral growth rates through general equilibrium mechanisms (Krueger et al. 1988; Timmer 1988; Stern 1989; Sah and Stiglitz 1984).

Direct agricultural pricing policies and the indirect effects of macroeconomic policies can in principle exert a significant long-run impact on economic structure. It is well known that many developing country governments have attempted to suppress and/or stabilize agricultural producer prices through procurement policies, export taxation, and/or export quotas. In addition, macroeconomic and industrial protection policies have indirect effects on agricultural prices: an industrial promotion policy is an implicit tax on activity in all other sectors (Corden 1974). It is widely observed that the indirect effects of trade and exchange rate policies designed to protect import-competing manufactures in developing economies embody a systematic bias against agriculture (Krueger *et al.* 1988).

In spite of this, current empirical investigations of the decline of agriculture do not incorporate measures of the effects of policies and conversely, quantitative analyses of development policy are typically content to address its impacts at the sectoral level rather than on overall economic structure. There is a need to reconcile the two strands of the literature.[4]

A second little-recognized source of structural change is the accumulation of human capital. It is now standard to include some measure of human capital in the aggregate production function (Lucas 1988; Schultz 1988; Barro 1991; Benhabib and Spiegel 1994; Lau *et al*. 1993; World Bank 1993); indeed, to omit human capital from an analysis of aggregate growth is to mis-specify the model. The same argument must apply to the analysis of structural change. Since agriculture and non-agriculture typically use skilled labor in differing proportions, it follows from Rybczinski that differences in the rate at which the skilled labor force grows must also influence the rate of agriculture's relative decline.

A third point regarding structural change concerns its variation over time. Poor countries typically tax agriculture relative to other tradable sectors, while industrialized countries tend to protect their farmers (Little, Scitovsky, and Scott 1970; Anderson and Hayami 1986; Lindert

1991, David and Huang 1996). A number of phenomena contribute to this change. Economic growth diminishes the importance of agricultural taxation as a source of government revenue; the desire to maintain agricultural exports motivates a reduction of export taxes; and finally, the political influence of agricultural producers diminishes less quickly than the decline of the sector's contribution to GDP. As a result, government policies are likely to increase the rate of agricultural decline in early stages of development, and reduce it subsequently.

Empirical motivation

In the econometric analysis that follows, we make use of data from Thailand, a middle-income developing economy. However, our arguments are based on three broad observations that are immediately recognizable in the development experiences of many developing countries. First, many developing-country governments tax agriculture not only directly, through pricing policies, but also indirectly, through macroeconomic and industrial policies aimed primarily at promoting industrialization (Krueger *et al.* 1988). Both direct pricing policies and indirect discrimination often result in suppression (and sometimes stabilization) of agricultural prices in the domestic market. For Thailand, Figure 1 shows how domestic agricultural prices relative to nonagricultural prices diverge from border prices—the prices that would have prevailed in the absence of the direct and indirect effects of government intervention. Domestic agricultural prices are both lower and less volatile than border prices. It follows that a part of what is usually taken to be a secular price trend is in fact due to the unaccounted effects of government policies.

<Figure 1 about here>

The second observation is that relative agricultural prices may decline in secular fashion, but policies affecting them may change over time as political economy dictates. For Thailand,

Figure 2 shows levels of the nominal protection rate for agriculture (a negative value indicates an effective tax on agriculture).[5] An inverse parabola fits these data very well,[6] and this highlights an additional empirical observation, that changes in the net taxation of agriculture are not necessarily linear.

<Figure 2 about here>

The third observation is that human capital accumulation is fundamental to the economics of agricultural decline. Quantitative estimates of the contribution of factor endowment growth to agricultural decline should include measures of growth in these endowments.

On the basis of these three observations, our empirical analysis takes the pioneering work of Martin and Warr (1993, 1994) as a starting point and extends it to incorporate agricultural pricing policies and human capital accumulation as additional factors explaining the relative decline of agriculture. Using a translog revenue function, we derive an expression for the share of agriculture in national income as a function of the border prices of commodities, agricultural pricing policies, and factor endowments including human capital. We then estimate the structural relationship between agriculture's GDP share and its determinants using a vector error-correction-mechanism (VECM) model. This approach is particularly versatile when dealing with endogeneity and nonstationary data in time series analysis. We estimate the model on Thai data by Johansen's (1988) method.

Our empirical results may be summarized as follows. First, agricultural price policies—effective agricultural taxes—are found to be the most important measures influencing the relative decline of agriculture in Thailand. Their relative contributions to the decline are greater than that of all other supply factors combined. Second, by increasing effective agricultural taxes, past Thai governments accelerated the rate of agricultural decline in the early stages of postwar economic development. Subseqent policy reforms in the late 1980s—a decade in which manufacturing income grew far faster than agricultural income—slowed the agricultural decline. Third, in contrast to Martin and Warr's finding that capital accumulation made the greatest contribution to the decline of Thai agriculture, we find renewed support for relative price trends as major contributing factors—inclusive, that is, of the effects of policies affecting the domestic sectoral terms of trade. Accumulation of physical and human capital are the second and third factors, after prices. Finally, sectoral asymmetries in technological progress play no significant role in the structural transformation.

3. The Model

In this section we present a model for estimating the determinants of agricultural decline for a time series of data from Thailand. The fundamental analytical construct is the revenue function (Dixit and Norman 1980; Woodland 1982). As we demonstrate, a translog flexible functional form of the revenue function yields tractable empirical relationships among variables of interest.

Consider a small open economy characterized by constant returns to scale and competitive market equilibrium. Assume that there are two aggregate outputs, agriculture (A) and non-agriculture (N), and four types of inputs: physical capital (K), labor (L), human capital (H), and land (R). The aggregate production possibility frontier can be defined by the implicit production possibility frontier

$$G(A, N, K, L, H, R) = 0 \tag{1}$$

For given factor endowments and technology, the allocation of resources between A and N—that is, the economically optimal position on the frontier—depends on relative commodity prices, and is found as the solution to maximization problem:

$$r(P, Z) = \max\left[PY : Y \in Y(Z)\right], \qquad P, Y \in \mathbb{R}^2, Z \in \mathbb{R}^4, \tag{2}$$

where r is maximized aggregate revenue or GDP; Y is the final goods vector, P is the vector of final goods prices, and $Y(Z)$ is the convex producible set given endowments $Z=\{L, K, H, R\}$.

The vector of output supplies $x(P, Z)$ is given by the gradient of the revenue function with respect to prices.

$$x_i(P,Z) = \partial r(P,Z) / \partial P_i \qquad \text{for } i = A, N. \tag{3}$$

We assume that the revenue function is twice continuously differentiable with respect to prices, so that the outputs in each sector are uniquely determined. As in standard international trade models, we also assume non-jointness in input quantities; that is, there exists a production function for each good. We also assume that the agricultural production is relatively less intensive in the use of both physical and human capital than is non-agriculture. For given technology and prices, Rybczinski's theorem then suggests that physical or human capital accumulation will reduce increase nonagricultural output and employment, and reduce agricultural output and employment.

Now consider the effects of technical change on the revenue function. Supposing (for expositional purposes) that technologies in both sectors are exogenously determined and are neutral with respect to factors, they can be represented by shift parameters in a vector T. We can then rewrite the revenue function as $r(P, Z, T)$. The corresponding supply functions are :

$$x_i = T_i f_i(P,Z) \quad \text{for } i = A, N. \tag{4}$$

In (4), an increase in T_i indicates Hicks-neutral technological change in sector i, and is equivalent to a price increase for the sector concerned, since an increase in T_i means higher production value for given factor use (Dixit and Norman 1980:137). Therefore, we can rewrite the revenue function as $r(TP, Z)$, where TP is a vector with elements $(T_A P_A, T_N P_N)$.

For estimation purposes, we approximate the revenue function by a translog functional form. Following Woodland (1982) and Kohli (1991), we can derive the translog revenue function from (1) as:

$$\begin{aligned} \ln r(TP, Z) = {} & \alpha_0 + \sum_i \alpha_i \ln(T_iP_i) + \frac{1}{2}\sum_i \sum_j \gamma_{ij} \ln(T_iP_i)\ln(T_jP_j) + \sum_k \beta_k \ln(F_k) + \\ & \frac{1}{2}\sum_k \sum_h \delta_{kh} \ln(F_k)\ln(F_h) + \sum_i \sum_k \phi_{ik} \ln(T_iP_i)\ln(F_k) \end{aligned} \qquad (5)$$

where ln denotes the natural logarithm;

r = total revenue;

P_i = prices of agriculture and non-agriculture ($i, j = A, N$);

F_k = quasi-fixed factor inputs ($k = K, H, L, R$); and

α_i, β_i, γ_{ij} are parameters to be estimated.

The revenue function is linearly homogeneous in prices, which implies the following restrictions on parameter values: $\sum_i \alpha_i = 1$; $\sum_i \gamma_{ij} = 0, \forall j$; and $\sum_i \phi_{ik} = 0, \forall k$.

Differentiating (5) with respect to $ln(TP_i)$ and imposing homogeneity restrictions yields sectoral share functions. The GDP share of sector i is a function of prices, factor supplies and technology parameters:

$$S_i = \alpha_i + \sum_j \gamma_{ij} \ln(T_jP_j) + \sum_k \phi_{ik} \ln(F_k), \ \forall j = A, N \text{ and } \forall k = K, L, H, R, \qquad (6)$$

where $S_i \equiv (P_iY_i)/\sum_j P_jY_j$ is the GDP share of sector i. By the properties of the revenue function, the output supply function and GDP shares are both homogenous of degree zero in prices. Thus we can normalize P_A by P_N. With constant returns to scale, we can express all quantities in intensive form, which we do by dividing all factor inputs by labor. Thus we have in one equation a complete description of the factors determining the GDP share of agriculture:

$$S_A = \alpha_A + \gamma_{AA} \ln\left(\frac{P_A}{P_N}\right) + \gamma_{AA} \ln\left(\frac{T_A}{T_N}\right) + \phi_{AK} \ln\left(\frac{K}{L}\right) + \phi_{AH} \ln\left(\frac{H}{L}\right) + \phi_{AR} \ln\left(\frac{R}{L}\right) \qquad (7)$$

Equation (7) can be viewed as representing the long-run equilibrium structure of the economy. The agricultural share depends on relative domestic prices, relative factor supplies, and differential rate of technical change. The model derivation provides a convenient vehicle for testing hypotheses concerning sources of decline in the agricultural share of GDP.[7]

In a small open economy with competitive markets and no interventions, the domestic producer price of a tradable agricultural product would be the same as the border price of an identical good in the world market, evaluated at the official nominal exchange rate and adjusted for transport, storage, and other costs. Observed relative domestic prices would thus encompass the true relative border price *and* the effects of taxes imposed by the government. Defining p as the observed domestic relative price, and letting p^* stand for the real border price, we would have

$$p = p^*(1 + \tau), \tag{8}$$

where τ is a measure of the total nominal protection rate, capturing both the direct effect of price interventions and the indirect effects of economy-wide on the domestic price (for a more detailed discussion, see Krueger *et al*. 1988). Empirically, we do not expect (8) to hold exactly, most obviously because some part of the intervention does not operate through market prices. Similarly, the impact of technical change on agriculture's GDP share need not be exactly equivalent to that of a change in domestic relative prices, since Hicks-neutrality is not assured. With these cautions in mind, we decompose the price term in (7) into its border price and policy components to obtain an estimable equation:

$$\begin{aligned} S_A = \beta_A + \beta_{Ap} \ln\left(p^*\right) + \beta_{A\tau} \ln\left(1 + \tau\right) + \beta_{AT} \ln\left(\frac{T_A}{T_N}\right) \\ + \beta_{AK} \ln\left(\frac{K}{L}\right) + \beta_{AH} \ln\left(\frac{H}{L}\right) + \beta_{AR} \ln\left(\frac{R}{L}\right) + \varepsilon. \end{aligned} \tag{9}$$

In (9), we allow for border prices, price policies, and technological change to have separate effects on the share of agriculture. In this equation ε is a stochastic error term indicating unobserved shocks and measurement errors.

Equation (9) incorporates both classical and non-classical factors in explaining agriculture's relative decline. In particular, it allows for the separation of an overall price effect into its components due to policies and border prices, and for Rybczinski effects from several types of factor endowment growth. The structural relationship in (9) provides a new framework for investigating long-run equilibrium relations and is feasible to be tested. We explore the details in the next section.

4. Estimation Issues and Hypothesis Testing

The long-run specification in (9) assumes free movement of factors among sectors and abstracts from adjustment processes. If the reallocation of factors among sectors in response to changes in technology, prices, and aggregate factor supplies is costly and/or occurs with a lag—as in most agricultural production— the aggregate share observed will reflect deviations from equilibrium.. It is thus important that the model capture dynamic responses and adjustment processes. We expect, moreover, to find that some of the explanatory variables in (9) are not strictly exogenous. This problem is more likely to occur if there is lagged transmission of the last-period errors into a current-period decision.

The conditional VECM form provides a convenient means to handle both the dynamic processes and the endogeneity problem in (9), and the relationship can be estimated by Johansen’s (1988) method. We can define the conditional VECM as follows:

$$\Delta Y_t = \Lambda_0 \Delta Z_t + \Gamma_1 \Delta X_{t-1} + \alpha\beta' X_{t-1} + \Phi D_t + \varepsilon_t \qquad (10)$$

where Y_t is a vector containing only endogenous variables; Z_t is a vector of weakly exogenous variables (included in the cointegration space); X_{t-1} is a vector of all variables; D_t is a vector of non-stochastic variables, such as a dummy variable for periods before and after a policy shift; and □ $_t$ is a stationary multivariate disturbance. For parameters to be estimated, □$_0$ and □$_1$ show the short run adjustments to changes in Z_t and $X_{t\text{-}1}$ respectively; □ represents the speed of adjustment, and □ is a matrix of long-run coefficients such that □□$X_{t\text{-}1}$ represents up to (n-1) cointegration relationships which ensure that variables in X_t converge to their long-run steady states. The linear combination of levels $X_{t\text{-}1}$ is just its combination which is stationary in levels, so □ $_t$ is I(0) and can be said to be white noise. The estimates of □□□show long run adjustment.

A test of the hypothesis that □$_{ij}$ is zero for j=1,...,r provides a test for weak exogeneity of the explanatory variables. The LR test statistic for this is T□log[(1–□*)/(1–□)], where □* are new eigenvalues for the restricted model. This statistic has a □^{2} distribution with $r(n\text{–}m)$ degrees of freedom, where r is the number of cointegration relationships, n the number of possible endogenous variables, and m the number of endogenous variables (Johansen and Juselius 1992).

In estimation, one important issue that must be addressed is that of a possible structural break in our time-series data or in the model relationship. The first possibility, that of a structural break in the data, can be easily checked when testing unit roots. The second can be addressed using the standard dummy variable technique and the conditional VECM form in model estimation.

The final estimation issue is to test the number of cointegration relationships. The null hypothesis is that there are at most r cointegration vectors and (n–r) unit roots as: H_o: □$_i$ = 0 for i = r+1,...,n; where only the first r eigenvalues are nonzero and □$_i$ is the eigenvalue. Two common likelihood ratio tests (the Trace and □-*max* statistics) can be used to test this hypothesis (Johansen and Juselius 1990).

According to the theory presented in Section 2, we expect the following *ceteris paribus* results. An increase in the real border price of agriculture would imply a greater share of agriculture in GDP, as would an increase in net agricultural protection (□) and an increase in the index of technical change in agriculture relative to that in non-agriculture. An increase in the physical capital-labor ratio should reduce the share of agriculture, as predicted by the Rybczynski theorem under the assumption that agricultural production is relatively less intensive than non-agriculture in the use of physical capital. By the same reasoning, an increase in the ratio of human capital to labor should also result in a lower share of agriculture.

5. Data Measurement and Trends

Measurement.

The analysis employs national-level data set based on official statistics covering 1951 to 1995.[8] The domestic relative price is calculated using implicit price deflators of agriculture relative to non-agriculture. Implicit price deflators are calculated by dividing output at current prices by output at a constant price. Real border prices represent relative commodity prices at the border when both direct and indirect pricing policies on agriculture prices are removed. These policies are measured as the total nominal protection rate (NPR). We calculated the aggregate NPR from the total NPR of the three major crops (rice, rubber, and maize), weighted by the share of each in total agricultural exports.[9] The aggregate NPR prior to 1960 is approximated by the percentage of agricultural export taxes to export values by crop. In the absence of direct data, the aggregate NPR between 1987 and 1990 is approximated by the average of the three earlier years. For the same reason, the NPR after 1990 is assumed to be constant at 15 per cent—a proximate number for the impact of indirect macroeconomic policies only. In this period, most direct agricultural trade policies had already been largely removed.

The total labor force is defined by the population aged over 13 years. Prior to 1960 we have data only for 1947 and 1954. We fill the gaps in the early part of the series by interpolation. The aggregate capital stock series is obtained from official sources for the period after 1970. For earlier years, we estimate the series recursively using the standard perpetual inventory method. The estimated depreciation rate is 5%, as suggested by Limskul (1988).

The measurement of human capital deserves careful attention since other parameter estimates might be sensitive to human capital measurement. In this paper we construct indices of human capital stock based on the returns to schooling. The basic idea is to combine a wage differential index (as a proxy for returns to schooling) with data on years of schooling of the labor force. This method is preferred to the more frequently used school enrollment data for two reasons. First, the gross enrollment ratio data—the number of enrolled children in the relevant age range divided by the total number of children in that age range when school is open—overestimate schooling investments and are likely to differ from net enrollment (Behrman 1993). Second, wage differentials should reflect differences in the marginal productivity of schooling. Thus the stock of human capital at year t, H_t, can be defined as a weighted sum over labor force subgroups defined by education:

$$H_t = \sum_{i=1}^{n} W_i S_i N_i ,$$

where W_i denotes the wage differential index; S_i the average years of schooling of members in the ith group; and N_i the number of workers in the ith group.

The land endowment is measured as cultivated land area. This measure takes account of increases in *effective* land area due to irrigation, which permits double-cropping. In the absence Technological change in non-agriculture is estimated by a total factor productivity growth (TFP) index, calculated using the growth accounting technique.[10]

Trends in Data.

Over the sample period, the domestic agriculture-non-agriculture price ratio has a slight but stable downward trend, falling by an average 0.45% per year (Figure 1). It appears to drift with persistent deviations from any underlying price level trend. In contrast, the real border price shows more variation and was much higher than the domestic ratio.

The calculated total NPR (Figure 2) shows a pronounced nonlinear trend. Between 1950 and 1975, a period of relatively high and increasing protectionism, the total NPR increased from around 20% to around 40% of the real border price. It then began to rise sharply from the early 1980s, following a significant relaxation of rice export restrictions in 1980-81. After 1983, the Thai government embarked on a liberalization program that resulted in significant reductions in non-agricultural trade barriers.

The agricultural share of GDP, measured at 1972 prices, is shown in Figure 3. It declined from over 50 % in 1951 to around 10 % in 1995, an average annual decline of 3.5 %.

<Figure 3 about here>

On the input side, all the series depicted in Figures 4 to 7 trended up or down for sustained periods. The capital-labor ratio (Figure 6) appears to be much less volatile than either agriculture's GDP share or relative prices. It increased during a period of massive public investment in infrastructure and communications in the early 1960s, then after a long period of slow growth in the 1970s, showed a new surge during the foreign investment-led boom of the 1980s.

<Figures 4-7 about here>

In Figure 7 we show the average human capital stock (H_1), and also, for comparison, the secondary school enrollment rate (H_2). At best, the enrollment ratio represents investment levels in human capital, and is likely to overestimate schooling investment.

The land-labor ratio (Figure 8) declined during the period covered by our data. However, it increased in 1960-1975, largely through expansion of cultivated area, which grew at an annual rate of 3.9 % during the 1970s then slowed to average of 1 % in the 1980s. The lower growth rate indicates the closing of the land frontier in the late 1970s. Finally, both sectoral indexes of technical change (Figure 9) show increasing trends, with that in non-agriculture outstripping that in agriculture after the late 1960s.

<Figures 8-9 about here>

6. Empirical Analysis

In this section we present two major results. First, we explore the long-run structural relationship between the agricultural share and its determinants, presenting estimates of the cointegrating vector from the conditional VECM model. Second, we present a decomposition of contributions to the relative decline of agriculture.

Long-run relationship between the share and its determinants

Before discussing the results from the conditional VECM model, we first report routine statistical tests on unit roots, cointegration and weak exogeneity, respectively.

Unit Root Test. We test for an $I(1)$ process in data, using the Augmented Dickey-Fuller (ADF) test with constant term and time-trend. Since the ADF test is sensitive to the order of augmentation, we use the Campbell and Perron (1991) criterion for choosing the optimal lag length in the ADF test to increase the test power. Too few lags in the ADF test may result in over-rejection of the null when it is true, while too many lags may reduce the power of the test. The non-standard critical values of the ADF test are from MacKinnon (1991). The test statistics

(Table 1) show no strong evidence against the null hypothesis of the unit roots of the variable of interest at the 5% significance level.

<Table 1 about here>

The results of unit root tests when allowing for a possible structural break are reported in Table 2. Overall, we again find no evidence against unit roots. Using the recursive minimum ADF *t*-statistics (column b in the table), we cannot reject the null even after allowing for the structural break in the series. Using our *a priori* information, we suspect a structural break associated with the major change in rice export policy in 1981 and the subsequent reduction in industry sector protection. We recalculate the ADF test, including dummy variables to capture changes in means (column c) and in trend (column d). The new results show similar results. We cannot reject the null with the possibility of the break in mean or in trend in all variables.

<Table 2 about here>

Cointegration Test The results of the unit root tests indicate that it is possible to define a stable relationship between the agricultural share and its determinants in terms of cointegration relationships. We employ Johansen's method to find the number of such relationships. In the Johansen model, we also include two exogenous dummy variables: *D*81 and *D*81**ln*(1+□). The *D*81 variable captures a change in the intercept after 1981, and the interaction term *D*81**ln*(1+□) captures a change in slope.

The Johansen test results, obtained using the CATS program (Hansen and Juselius 1995), are reported in Table 3. The first null hypothesis is that there is no cointegrating vector (Ho: r= 0) with the alternative is that there is one cointegrating vector (Ha: p-r). The second null is that there is at most one cointegrating vector against two, and so on. We report the trace statistic, its critical value at the 10% significance level, and estimated eigenvalues. The trace statistics indicate that there are three cointegrating relationships among variables at the 10% significance level. Since there is no clear economic explanation for these extra relationships, we impose the restriction that there is one cointegration relationship in our model. In doing so we follow current literature suggesting emphasis on the economic relationship rather than statistical relationship (Harris 1995). The implication of this choice is some efficiency loss in estimation.

<Table 3 about here>

Table 4 shows the normalized estimates of β and α obtained from the Johansen test. The β vector sets the element associated with the share equal to 1, representing the cointegration relationship. The α vector represents the speed of adjustment for each endogenous variable from the last period disequilibrium.

<Table 4 about here>

Weak Exogeneity Test. The low *t*-statistics on α suggest that human capital, technology, and the real border price might be weakly exogenous.[11] We conduct a weak exogeneity test for these variables. The test statistic of the null of $\alpha_2 = \alpha_4 = \alpha_6 = 0$ is the Likelihood Ratio (LR) test. The LR statistics shown at the end of the table do not exceed the Chi-square with 3 degrees of freedom, which is 6.25 at the 10 % level. Thus we do not reject the null hypothesis that these variables are weakly exogenous.

Knowing that some variables are weakly exogenous, we reapply Johansen's method in estimating Eq. (10)—the conditional VECM form. Endogenous variables in this construct are the share, the effective agricultural tax, the capital-labor ratio, and the land-labor ratio. Weakly exogenous variables are the real border prices, the technological bias, and the average human capital. Table 5 shows results of parameter estimates of the multivariate system in Eq. (10).

<Table 5 about here>

The diagnostic tests of the model are shown at the end of Table 5.[12] For goodness of fit of the model, the Ljung-Box test shows no serial correlation problem of the first 10 lags; and the Shenton-Bowman test, $\chi^2(8)$, does not indicate non-normality of multivariate residuals. For each equation, the Lagrange Multiplier (LM) test of autoregressive conditional heteroscedasticity (ARCH) of order 1 in the residuals shows no sign of a problem, and the Shenton-Bowman test does not show any non-normality problem in the residuals.

Estimates from the conditional VECM model. From Table 4, estimates of the cointegration vector β, or the long-run structural relationship between the share and its determinants, can be represented as follows:

$$
\begin{aligned}
S_A = {} & 1.29 + 0.10\,ln(p^*) - 0.42\,ln(1+\tau) - 0.96\,ln\left(\frac{K}{L}\right) + 0.26\,ln\left(\frac{R}{L}\right) \\
& - 0.05\,ln\left(\frac{H}{L}\right) + 0.03\,ln\left(\frac{T_A}{T_N}\right)
\end{aligned}
\tag{11}
$$

The coefficients on all of the variables have the predicted signs. A one per cent increase in the real border prices will increase the agricultural share of total GDP by 0.1%. Agricultural pricing policies—effective agricultural taxes—have a negative effect on the share. Increasing the effective agricultural tax rate by 1 per cent will decrease the share of agriculture by 0.42%. This impact is much larger than the impact of the real border prices.

A one per cent increase in the capital stock per unit of labor will reduce agriculture's share by approximately 0.96%. This is consistent with the Rybczynski theorem and our expectations about the relative capital-intensity of sectors. Given prices, technology, and the endowments of labor and other factors, accumulation of physical capital will increase factor productivity in non-agriculture, a process which will result in the withdrawal of complementary factors (unskilled labor) out of agriculture. In this *ceteris paribus* example, agricultural output would fall in absolute terms as a result; in practice, with all factor endowments growing but the capital stock at a faster rate, the growth rate of agriculture lags behind that of the economy as a whole.

An increase in human capital has the same type of impact. A one per cent increase in average human capital will decrease the share of agriculture by 0.05%. Conversely, a one per cent increase in the land-labor ratio has a positive impact on agriculture's share. This result is once again consistent with the Rybczynski theorem, since agriculture is land-intensive relative to other sectors. Finally, a one per cent increase in the index of differential technical change (agriculture relative to non-agriculture) increases the share of agriculture by 0.03%.

Decomposition of the factor contribution.

Using the estimates from the cointegration regression, we can decompose the factors contributing to long run changes in the agricultural share in Thailand. The results are shown in Table 6, for the entire period and for the subperiods before and after the beginning of the reform phase. For each period we report the average rate of change in each explanatory variable and its contribution to the overall decline in the share of agriculture, calculated as 100 times the average annual change of each variable multiplied by the relevant coefficient from Eq. (11).

<Table 6 about here>

Consider first the results for the whole period. Movement in the real border price accounts for only 7.2 % of the decline in agriculture’s share. Sharp increases in the capital-labor ratio and average human capital also contribute to the decline, by 67.8 and 22.2 %, respectively. On average, agricultural pricing policies show a benign impact on Thai farmers over almost five decades, a finding that is both counterintuitive and at odds with the historical evidence.

By allowing for a structural break, however, we reach conclusions that contrast strongly with those for the model with no break. The last four columns of Table 6 show the relative contribution of each factor before and after 1980, respectively. Each variable's contribution is calculated using the coefficients and average annual changes corresponding to that sub-period. These results indicate that agricultural pricing policies were the most important factors contributing to the decline during 1951-1980. The relative impact of effective taxes is greater than that of all the other supply factors combined. Compared with other factors, real border price changes helped offset the pressure from policy in this period.

During 1981-1995, the increase in the capital-labor ratio made the most important contribution to agricultural decline, followed by real border price trends and growth in the land-labor ratio. This result is consistent with the historical record: during this period, capital accumulation was at its highest rate in decades; agricultural prices in world markets were lower than they had been during the previous three decades; and land availability was diminished by the closing of the land frontier. Most importantly, policies were the only sources of an *increase* in agriculture's GDP share during this period.

Finally, in both periods, differences in technological change between sectors made an insignificant contribution to the decline of agriculture.

It is worth comparing these results with those of Martin and Warr.[13] Their studies concluded that demand-side factors, operating through relative commodity prices, are less important than supply-side influences, which are largely determined by capital accumulation. For Thailand, they found that changes in the capital-labor ratio contributed 72% of the total decline in the agriculture share in 1960-85, and relative price effects contributed the remainder.

Our analysis, taking government policies and policy trends as well as human capital accumulation into consideration, suggests a reinterpretation both of the conventional wisdom of the development economics literature and of the earlier empirical results. First, relative rates of factor accumulation are clearly important, a finding we share with Martin and Warr. However, once the model is refined to include a more disaggregated definition of "capital", we find proportionally different contributions from physical and human capital accumulation. Moreover, our results show the relative contribution of each form of capital to alter over time.

Second, while a model fitted to the entire series confirms the finding that demand side (price) influences are less important than factor endowment growth, separate identification of policy trends and the division of the data into sub-periods once again reveals a generally larger price effect story and a major role for policy. Offsetting border price and price policy trends may have resulted in an underestimate of price effects in earloier studies. Ultimately, however, the *ranking* of influences over agriculture's decline in our study concurs with that of Martin and Warr, and not with the "conventional" view: Engel effects and sectoral differences in technical progress are in the main, secondary to the process of structural change.

7. Concluding Remarks

In this paper we define the long run structural relationship between the agricultural share and its determinants in the form of a cointegration relationship, derived from an underlying economic model. The vector error-correction-mechanism (VECM) form applied to this structural relationship captures both the endogeneity of variables and their dynamic adjustments. The Johansen reduced rank regression was used to estimate this model.

Our study shows that agricultural pricing policies, which are equivalent to effective agricultural taxes, have played the dominant role in agricultural decline at the early stages of development and in offsetting the decline in later stages. Physical and human capital accumulation have also been important, yet still secondary to pricing policies. Engel effects and differential rates of technical change between sectors have had relatively small impacts. Our results indicate that ignoring the influence of policy and the effects of structural change can lead to erroneous conclusions.

Our findings have important policy implications. Despite the secular decline of agriculture during the course of economic development, government policy is not neutral. Development policy, by taxing agriculture, accelerates its decline. During the early development period, effective agricultural taxes are important “push” factors for agricultural decline. The elimination of these taxes (as in the later period in Thailand) can be a powerful instrument in slowing agricultural decline. More importantly, effective taxes on agriculture appear to offset price trends, an interesting issue for further research.

These findings enrich the economic development literature by clarifying the relative importance of sources of the major change in economic structure that occurs during development. While our results rely on data from Thailand, the ubiquity of direct and indirect agricultural pricing policies as expressions of economic development strategy suggests that the same analytical framework applied to other developing countries might be expected to produce similar findings. The nature and the quality of development policies that contribute to sectoral resource reallocation, migration, and urbanization in developing countries may have a significant impact on long-run welfare growth. One implication is that the previous emphasis on "getting agricultural prices right", which has dominated much policy-oriented research in the past, may be warranted.

Our findings indicate several directions for future research. One aspect of effective agricultural taxes not explored here is their role in stabilizing domestic prices. Policy recommendations to eliminate effective agricultural taxes, if implemented, will not only lead to higher domestic producer prices, but may also increase variances. Measuring the costs and benefits of price stabilization is thus one topic for future research. Another is a more detailed exploration of the political economy of changes in agricultural policies, from discrimination to protection. Modeling agricultural protection as an endogenous outcome of economic development processes could be productive. Lastly, regarding econometric methods, modeling nonlinear relationships in the cointegration vector should provide another means to improve the quality of estimates. However, the theoretical and statistical foundations of this technique are still in their infancy.

References

Anderson, K., 1987. On why agriculture declines with economic growth. Agricultural Economics 1, 195-207.

Anderson, K. and Y. Hayami, 1986. The Political Economy of Agricultural Protection: East Asia in International Perspective. Allen and Urwin, Sydney.

Behrman, J. R., 1993. Essay on selected issues related to the contribution of human capital to economic development. Mimeo.

Banerjee, A. et al., 1993. Integration, Cointegration, and Error Correction in Econometrics. Oxford University Press, London.

Barro, R. J., 1991. Economic growth in a cross section of countries. Quarterly Journal of Economics 106, 407-443.

Benhabib, J. Spiegel, M. M., 1994. The role of human capital in economic development: evidence from aggregate cross-country data. Journal of Monetary Economics 34, 143-173.

Campbell, J.Y. and Perron, P., 1991. Pitfalls and opportunities: what macroeconomists should know about unit roots, in Blanchard, O.J. and Fischer, S. (Eds), NBER Macroeconomics Annual 1991. MIT Press, Cambridge, MA.

Corden, W. M., 1974 Trade Policy and Economic Welfare. Oxford University Press, London.

David, C.C., and Huang, J., 1996. Political economy of rice price protection in Asia. Economic Development and Cultural Change 44(3), 463-484.

Dickey, D. A. and Fuller, W. A., 1981. Likelihood ratio statistics for autoregressive time series with a unit root. Econometrica 49, 1057-72.

Dixit, A. and Norman, V., 1980 The Theory of International Trade. Cambridge University Press, Cambridge, UK.

Ethier, W., 1988. Higher dimensional issues in trade theory, in R. W. Jones and P. B.Kenen (Eds), Handbook of International Economics. North-Holland, Amsterdam, pp. 131-184.

Hansen, H. and Juselius, K., 1995 CATS in RATS: cointegration analysis of time series.

Harris, R. I. D., 1995. Using Cointegration Analysis in Econometrics Modeling. Prentice-Hall. Englewood Cliffs, NJ.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 231-54.

Johansen, S. and Juselius, K., 1992. Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for UK. Journal of Econometrics 53, 211-44.

Johnston, B. F., and J. W. Mellor, 1961. The role of agriculture in economic development. American Economic Review 51(4), 566-593.

Krueger, A., M. Schiff, and A. Valdes. 1988. Agricultural incentives in developing countries: measuring the effect of sectoral and economy wide policies. World Bank Economic Review 2, 255-271.

Kohli, U., 1991. Technology, duality, and foreign trade: The GNP function approach to modeling imports and exports. University of Michigan Press, Ann Arbor.

Lau, Lawrence J. et al., 1993. Education and economic growth. Journal of Development Economics 41, 45-70.

Limskul, K., 1988. The sectoral capital stock, employment, and sources of economic growth in Thailand. Paper no. 40, Economic Research Center, Faculty of Economics, University of Nagoya, Japan.

Lindert, P. H., 1991. Historical patterns of agricultural policy, in C. Peter Timmer (Ed), Agriculture and The State: Growth, Employment, and Poverty in Developing Countries. Cornell University Press, Ithaca, pp. 29-83.

Lipton, M., 1977. Why Poor People Stay Poor: Urban Bias and World Development. Temple Smith, London.

Little, I, Scitovsky, T. and Scott M., 1970 Industry and trade in some developing countries, Oxford University Press, London.

Lewis, W. A., 1954 Economic development with unlimited supplies of labor. Manchester School, 139-191.

Lucas, R. 1988. On the mechanics of economic development, Journal of Monetary Economics, 22, 3-22.

Mackinnon, J., 1991 Critical values for cointegration tests. In R.F. Engle and C.W.J. Granger, (Eds), Long-run Economic Relationships. Oxford University Press, London, pp. 267-276.

Martin, W. and Warr, P.G., 1994. Determinants of agriculture's relative decline: Thailand. Agricultural Economics 11, 219-235.

Martin, W. and Warr, P.G., 1993. Explaining the relative decline of agriculture: a supply-side analysis for Indonesia. World Bank Economic Review 7(3), 381-401.

Perron, P., 1990. Testing for a unit root in a time series with a changing mean. Journal of Business and Economic Statistics 8, 153-62.

Punyasavatsut, C., 1998 Economic growth and structural change in Thailand, 1950-1995. Unpublished Ph.D. Dissertation, University of Wisconsin-Madison.

Sah R. K., and Stiglitz, J.E., 1984. The economics of price scissors. American Economic Review 74, 124-138.

Schultz, T. W., 1953. The Economic Organization of Agriculture. McGraw-Hill, New York.

Schultz, T. W., 1988. Education investments and returns. In Hollis Chenery and T. N. Srinivasan (Eds.), Handbook of Development Economics. North Holland, Amsterdam.

Stern, N., 1989. The economics of development: a survey, Economic Journal 99, 597-685.

Timmer, P., 1988. The Agricultural Transformation. In Chenery, H. and Srinivasan, T. N., (Eds.), Handbook of Development Economics, vol. 1. North-Holland, Amsterdam

Woodland, A. D., 1982. International Trade and Resource Allocation. North-Holland, Amsterdam.

World Bank, 1993. The East Asian Miracle: Economic Growth and Public Policy. Oxford University Press, New York.

Table 1. Augmented Dickey-Fuller Test Statistics for Unit Roots, 1951-95 (With a constant term and a time trend variable).

$$\Delta X_t = \beta_0 + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

Variable	ADF Test Statistics	Lag-length	Critical Value at 5%
Share of agriculture	-3.3752	3	-3.5217
Relative border price	-2.4548	1	-3.5162
Tax rate	-3.3438	0	-3.5136
Capital-labor ratio	-1.4823	1	-3.5162
Land-labor ratio	-0.8279	2	-3.5189
Average Human capital	-1.6061	0	-3.5136
Differential technical changes	-1.8283	1	-3.5162

Note: Lag lengths are set by Campbell and Perron (1991) criterion.

Table 2. Recursive and Standard Augmented-Dickey Fuller tests of unit roots, 1951-1995.

$$\Delta X_t = \alpha_0 + \alpha_1 X_{t-1} + \alpha_2 t + \sum_{j=1}^{p} \beta_j \Delta X_{t-j} + \alpha_3 D + \varepsilon_t$$

where: (1) in the shift in trend model: $D = t$ for $t > k$, and D = 0 , otherwise.
(2) in the shift in mean model: $D = 1$ for $t > k$, and D = 0, otherwise.
Assume that the break period, k, is 1981.

	Recursive Statistics		ADF-test statistics	
	Mean- Statistics	Min- statistics	(assuming 1981 as date of break)	
	(a)	(b)	Mean-shift (c)	Trend-shift (d)
ln(p*)	-1.224	-3.038	-1.899	-2.006
ln(1+□)	-0.994	-3.126	-2.882	-3.065
ln(K/L)	-0.923	-2.070	-1.851	-2.276
ln(Aa/Ana)	-1.104	-2.190	-1.055	-1.673
ln(R/L)	-0.866	-2.294	-1.632	-0.160
ln(H/L)	-0.792	-1.914	-1.672	-1.615

Note: The critical value at 10% is –4.00 (Banerjee et al. 1992, Table1). The critical value at 10 % is –3.13.

Table 3. Johansen test for the number of cointegration relationships

Ho	Ha	Eigenvalue	Trace Statistics	Critical value at 5%	Note
0	7	0.640118	147.2360	124.2	*
1	6	0.536959	103.2908	94.15	*
2	5	0.477138	70.18338	68.52	*
3	4	0.389483	42.30052	47.21	
4	3	0.251439	21.08222	29.68	
5	2	0.173869	8.629287	15.41	
6	1	0.009633	0.416216	3.76	

* denotes rejection of the hypothesis at 5% significance level

Note: Calculation is based on the model with linear deterministic trend in data.

Table 4. Results of the Johansen Estimates of the VECM (based on one cointegrating regression, r=1)

	S_A	ln(p*)	ln(K/L)	ln(H/L)	ln(R/L)	ln(T_A/T_{NA})	ln(1+□)
□ = [	1.000	-0.10	0.96	0.05	-0.26	0.03	0.42
□'= [	-0.173	-1.161	0.07	0.036	0.209	0.115	-1.671
t-value	(-2.115)	(-1.187)	(3.119)	(0.199)	(2.397)	(0.427)	(-4.744)

<u>Test Restriction on $□_2=□_4=□_6$:</u> The Likelihood Ratio (LR) test, $□^2(2) = 4.05$

Table 5. Results of Johansen Estimates of Equation (10) and Diagnostic Tests.

	D(Share)	D(Capital-labor Ratio)	D(Land-labor Ratio)	D(Tax)
D(H/L)	-0.061	0.051	-0.026	0.139
	(-0.924)	(3.196)	-0.025	(0.531)
D(Aa/Ana)	-0.057	-0.012	0.19	0.074
	(-1.287)	(-1.146)	(2.5)	(0.424)
D(p*)	0.032	0.009	0.052	0.403
	(1.463)	(1.647)	(1.402)	(4.713)
D81	-0.014	0.026	-0.2	0.034
	(-0.771)	(6.021)	(-2.853)	(1.099)
D81*tax	0.009	-0.038	0.287	-0.071
	(0.273)	(-4.766)	(2.182)	(-1.241)
Speed of Adjustment	-0.176	0.094	0.326	-1.218
	(-2.04)	(4.519)	(2.205)	(-3.584)
<u>Diagnostic Tests</u>				
LM Test for ARCH(1)	0.804	1.147	2.125	0.586
S-B Test for normality	2.070	1.641	1.441	0.819

<u>Goodness of Fit</u>
Test for Autocorrelation: Ljung-Box (10), χ^2 (156) = 178.124, p-value = 0.07.
Normality test: Shenton-Bowman χ^2 (8) = 34.143, p-value = 0.00.

Note: D stands for first difference.
Value in parenthesis is t-statistics.
All variables are in logarithmic terms, except the share.

Table 6. Decomposition of the Decline in Agriculture's Share in GDP, Thailand 1951-1995.

	Whole Period		1951-1980		1981-1995	
	Avg. Rate of Change	Percent Contribution	Avg. Rate of Change	Percent Contribution	Avg. Rate of Change	Percent Contribution
Price Effects:						
Real Border Prices	-0.0048	7.18	0.0117	-12.82	-0.0358	137.41
Tax Equivalence	-0.0032	-19.65	0.0138	62.59	-0.0340	-542.75
Non-Price Effects:						
Capital-Labor Ratio	0.0048	67.85	0.0035	35.80	0.0074	268.78
Land-Labor Ratio	-0.0050	19.33	-0.0004	1.05	-0.0138	136.58
Average Human Capital	0.0322	22.25	0.2003	10.33	0.0544	97.41
Differential Rates of Technical Changes b/w Sectors	-0.0067	3.05	-0.0091	3.06	-0.0022	2.57
Total		100		100		100

Notes

---

[1] The early literature also identifies institutional rigidities creating inefficiency in agricultural factor use as a source of agricultural decline (Lewis 1954; Johnston and Mellor 1961). Lacking empirical support, these theories have languished.

[2] More curiously still, the "classical" literature reveals no agreement on whether agricultural factor productivity is can be expected to have grown more or less rapidly than factor productivity in non-agricultural sectors. Early structural change theories maintained that technical progress, occurring more slowly in agriculture than in other sectors, would enhance demand-side effects by causing resources to flow out from agriculture to sectors where their marginal productivity was higher (e.g., Johnston and Mellor 1961; Schultz 1953). Some later theorizing suggested instead that faster relative technical progress rates in agriculture would cause the sector to contract, as the resulting outward shift of the agricultural supply function against inelastic demand would turn the domestic terms of trade against the sector (Timmer 1988). The differences between these two stories are not limited to assumptions about relative rates of technical progress. The first implicitly assumes that the domestic terms of trade will not be influenced by sectoral productivity growth differences. The second implicitly assumes that they will.

[3] In the case of many goods and many factors, the theorem holds in an average sense (Ethier 1988).

[4] This link has not escaped the attention of developing country policy makers. Beginning in the 1920s with Preobrazhensky's "price scissors" industrialization policies in the USSR, governments have exploited their influence over the domestic sectoral terms of trade in efforts to extract agricultural "surplus" (Lipton 1977).

[5] For the calculation of the nominal protection rate see Krueger et al. (1988).

[6] Excluding the apparently anomalous first observation, a curve fitted by MS Excel gives $y = 0.0004x^2 - 0.0185x - 0.169$, with R^2=0.774. Retaining the first observation reduces the R^2 to 0.51.

[7] It is important to note that K, H, R and L represent aggregate endowments not sectoral factor allocations.

[8] Details of data sources can be obtained in Punyasavatsut (1998).

[9] Lack of data preclude a more general approach. The three crops identified account for an average of about 50% of agricultural production by value during the period (Siamwalla and Setboonsarng 1990).

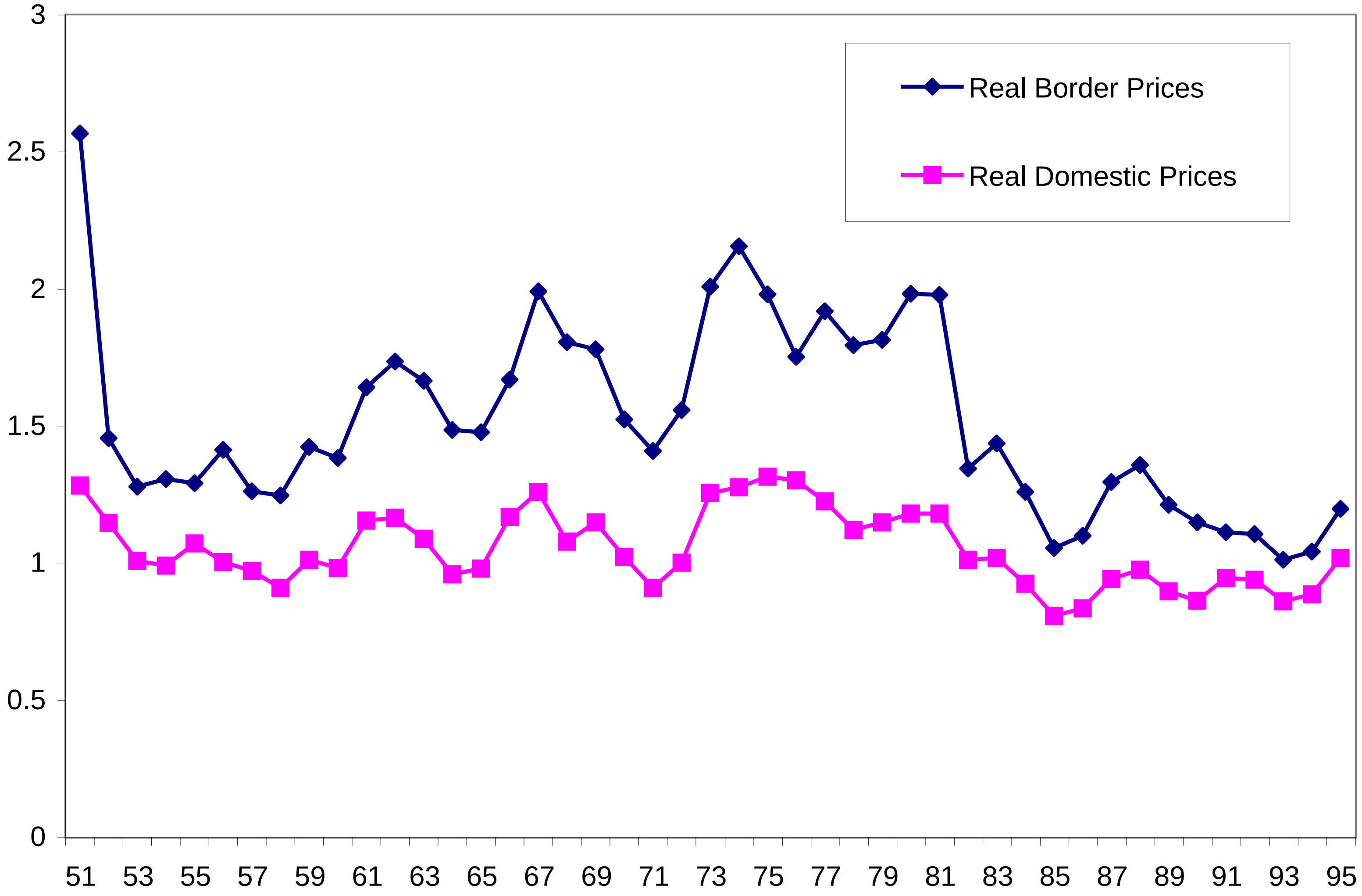
[10] Our definition of the technical progress ratio thus differs from previous studies such as Martin and Warr (1994), who used a time trend to capture the differential rate of technical change between agriculture and the rest of the economy. A time trend shows a very strong correlation with other upward-trending data such as the capital-labor ratio and human capital to labor ratio.

[11] Weak exogeneity of human capital, technology and the real border prices means that these variables contain no information about the long-run □ (a matrix of long-run coefficients) since the cointegration relationships do not enter into these equations. Therefore the estimation model can condition on the weakly exogenous variables for greater efficiency in estimation (Harris 1995).

[12] All estimates and diagnostic test results are obtained using RATS software.

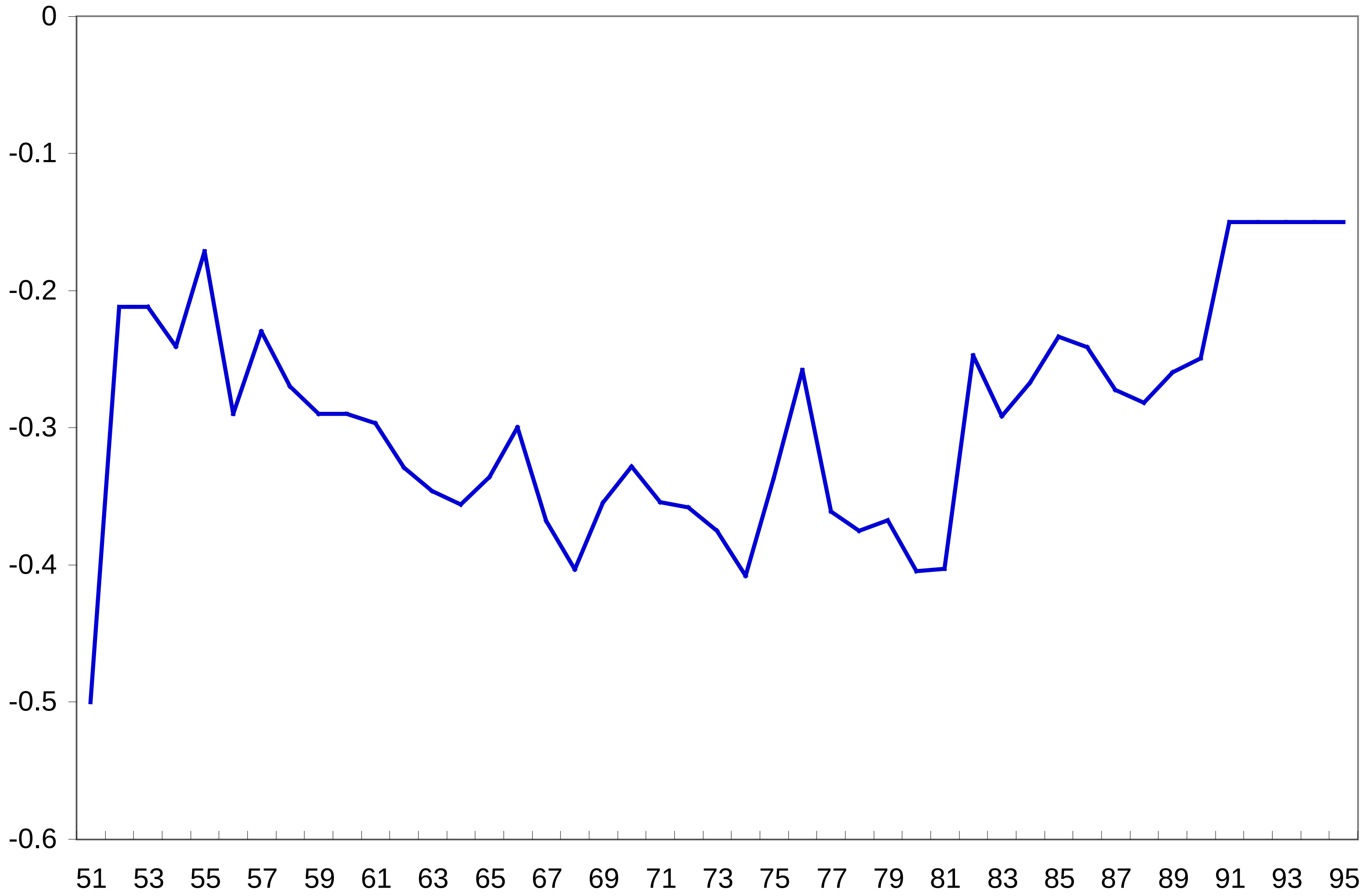
[13] Models used in the Martin and Warr studies differ slightly from that presented here. For their Thailand study, only physical capital and labor inputs included in the production function. Two relative price terms—agriculture relative to non-traded goods, and manufactures relative to non-traded goods—were used. The econometric relationship was modeled as a system of equations with generalized autoregressive distributed lags. The endogeneity problem was resolved by using two-stage least-squares.

Figure 1: Real Domestic and Border Prices, Thailand



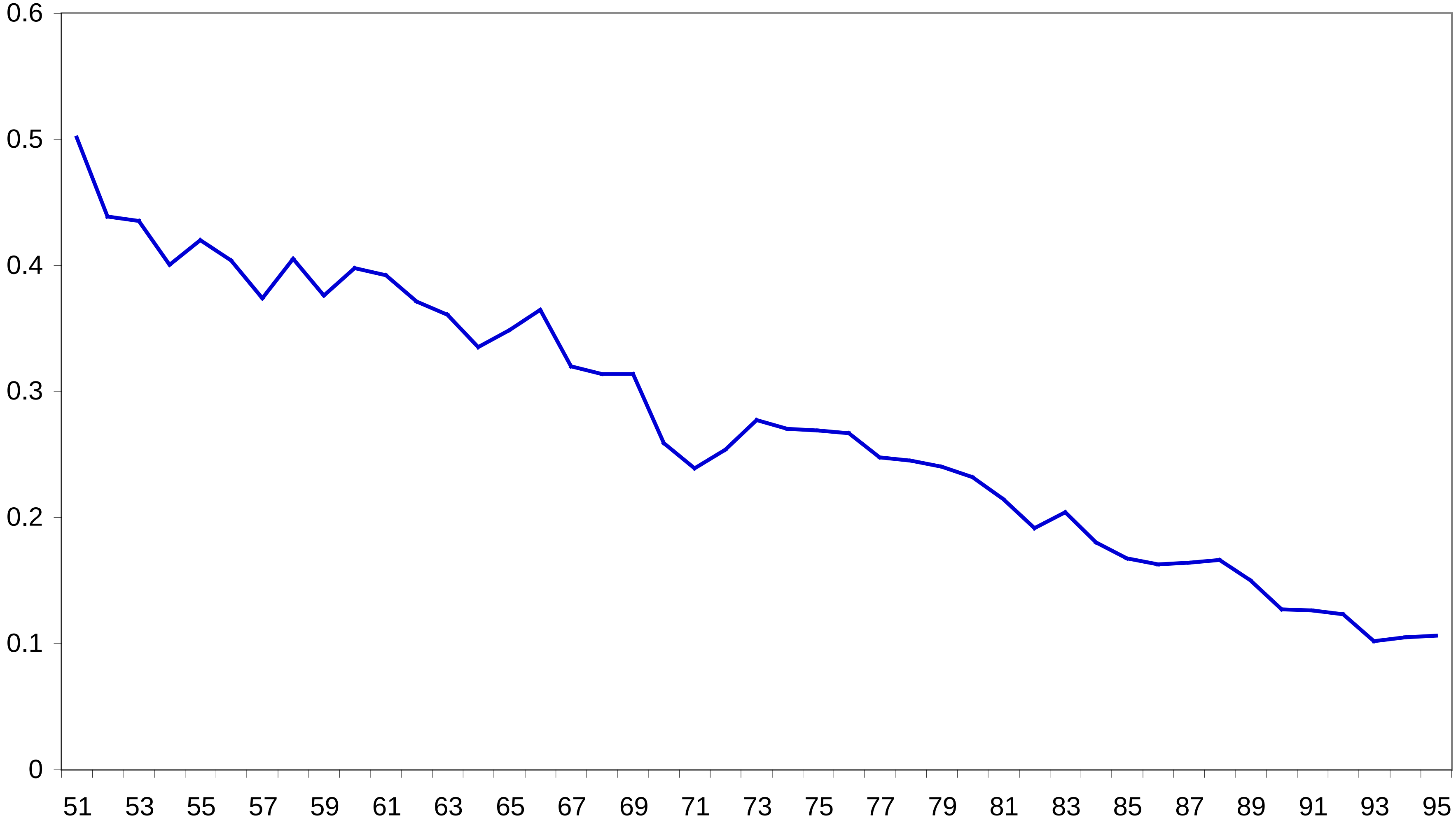
Source: NESDB and Chaiyuth 1998

Figure 2: Nominal Protection Rate, Thailand



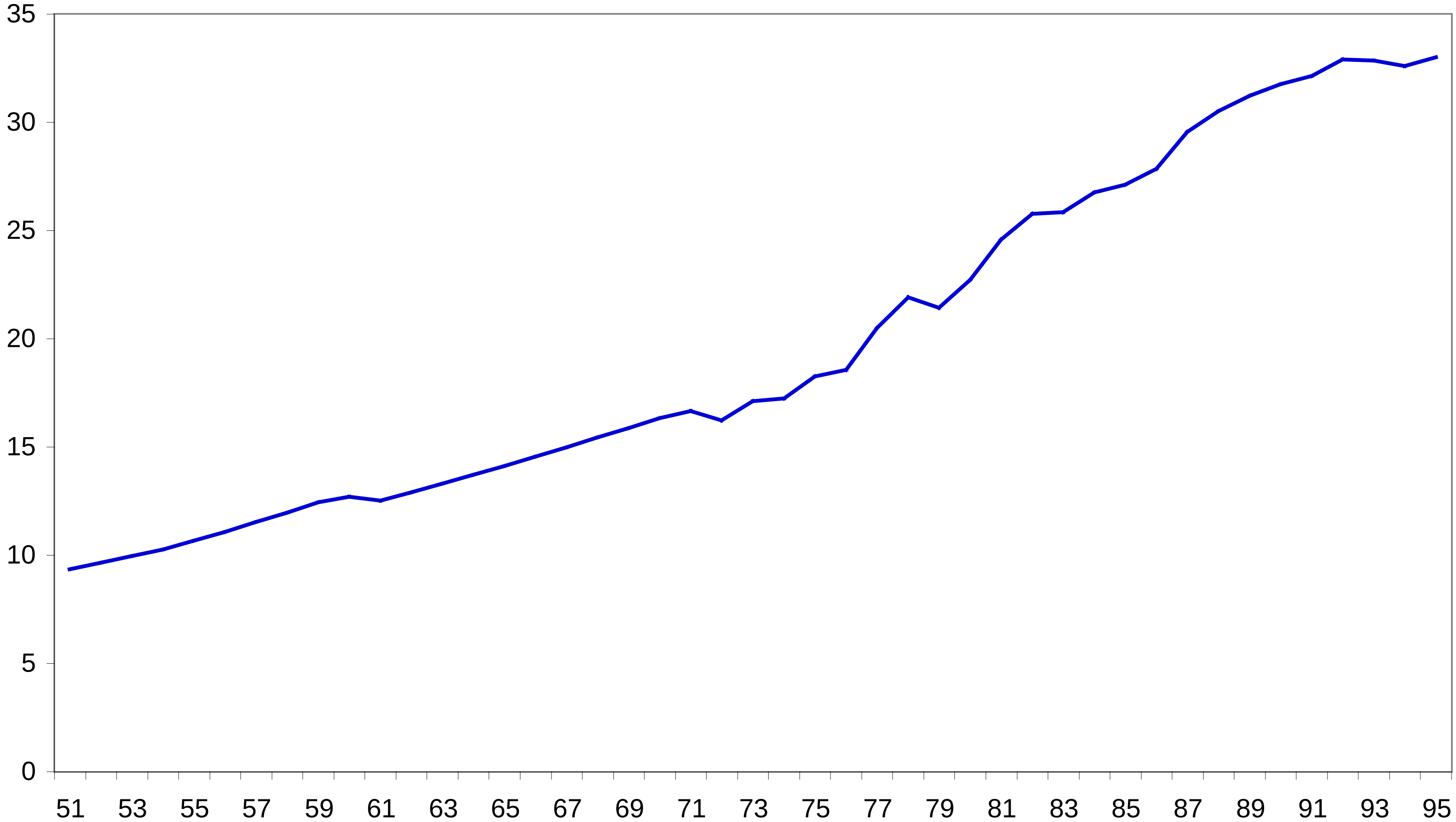
Source: Siamwalla and Setboonsarng 1988, and Chaiyuth 1998

Figure 3: GDP Share of Agriculture, Thailand



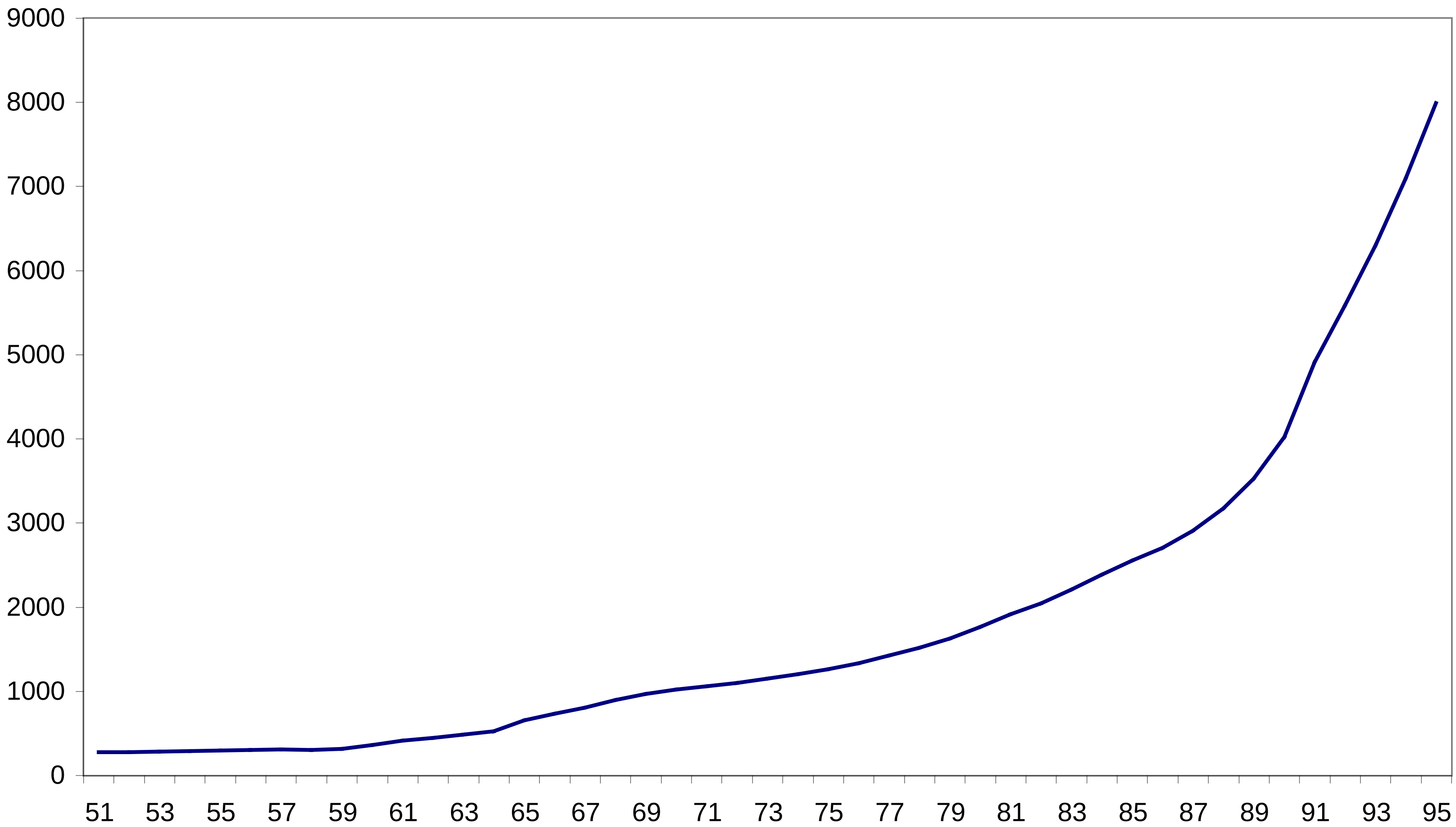
Source: NESDB

Figure 4: Labor Force (millions), Thailand



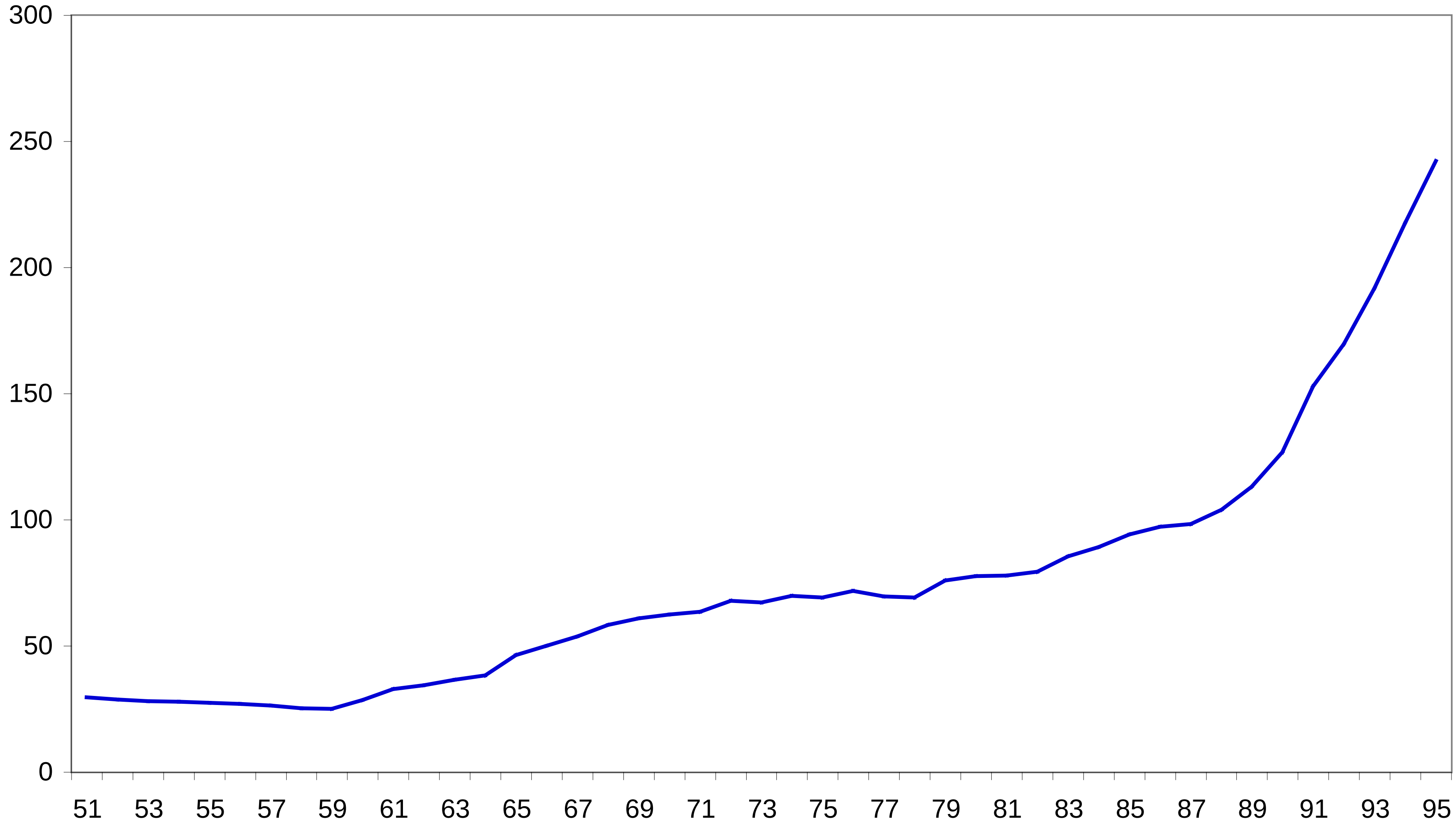
Source: Labour Survey Statistics, various years

Figure 5: Net Capital Stock (billions of baht),Thailand



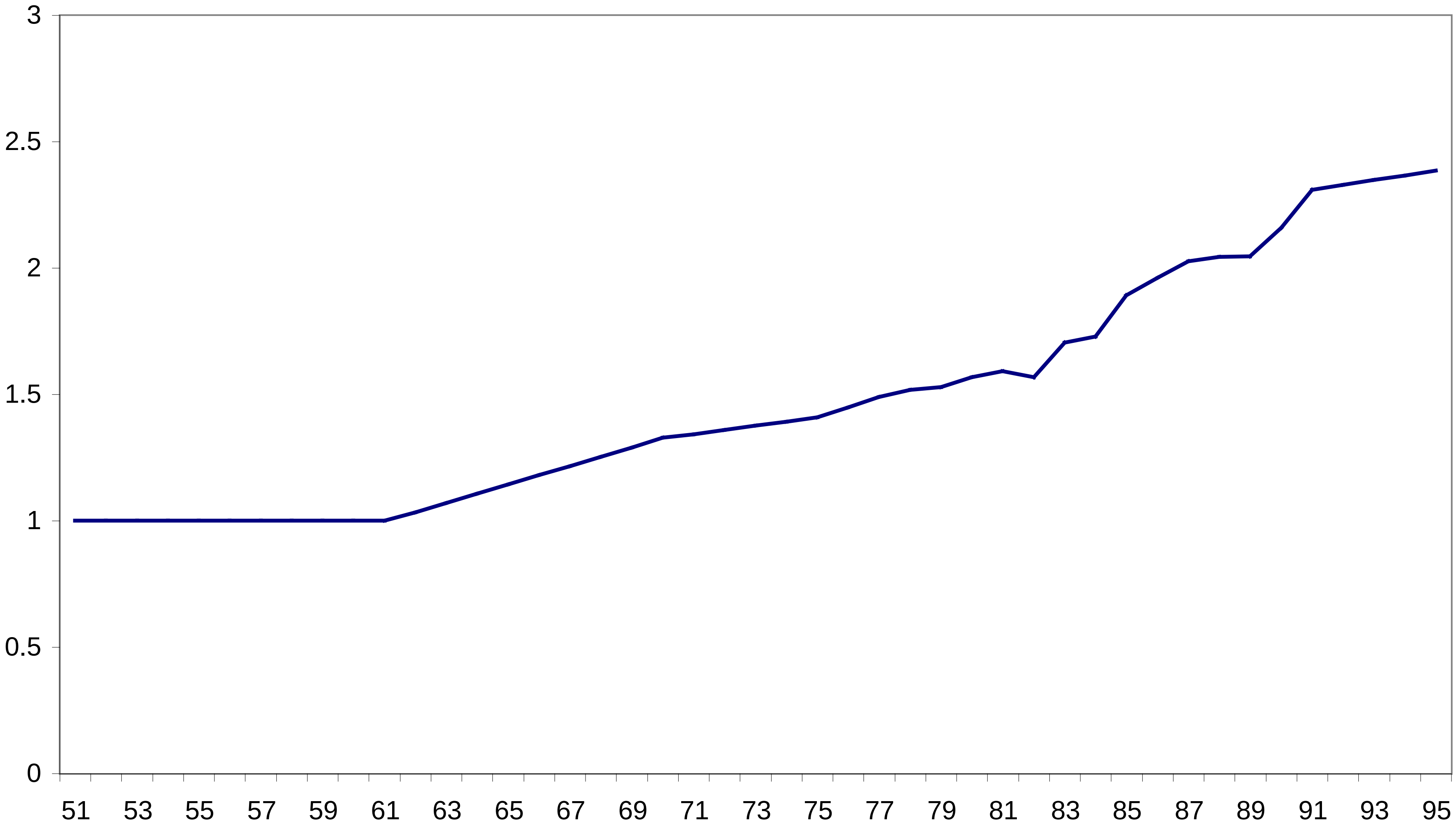
Source: NESDB and Chaiyuth (1998)

Figure 6: Capital-Labor Ratio (thousands of baht per worker), Thailand



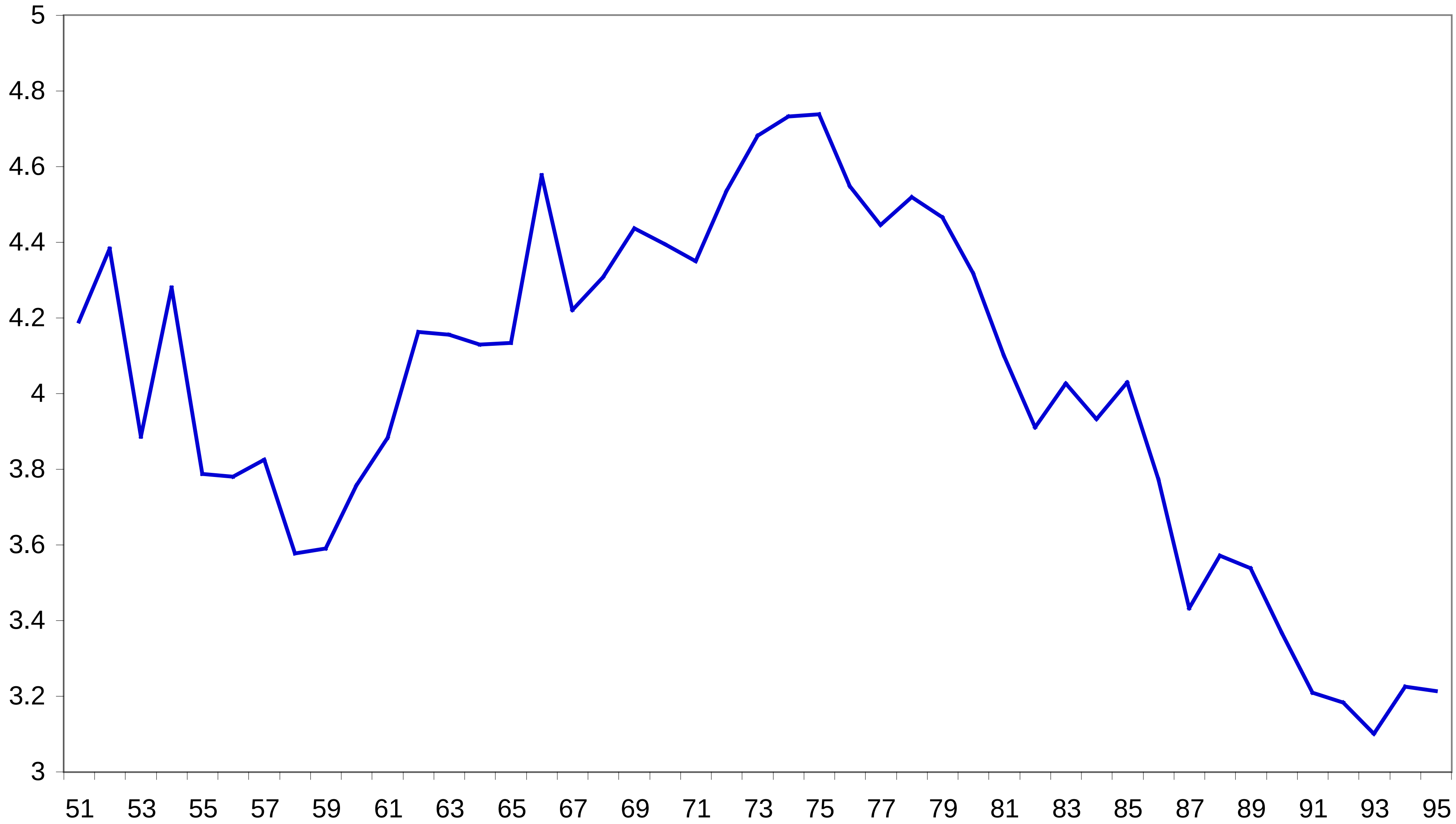
Source: As for Figures 4 and 5

Figure 7: Human Capital Index, Thailand



Source: Labor Force Survey Statistics and Chaiyuth 1998

Figure 8: Agricultural Land Per Worker (Rai), Thailand



Source: Office of Agricultural Economics

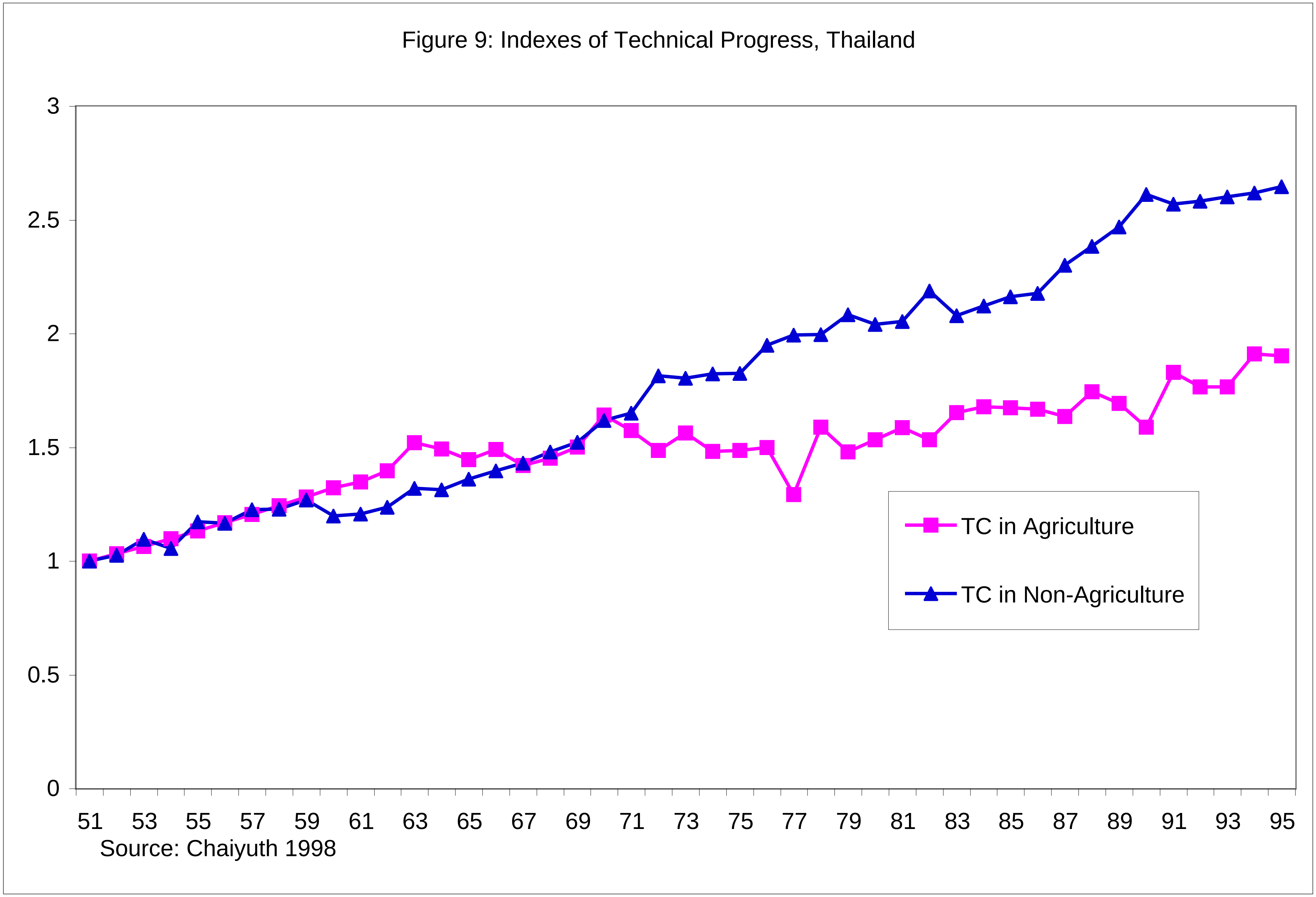


Figure 9: Indexes of Technical Progress, Thailand
3
2.5
2
1.5
1
0.5
0
51
53
55
57
59
61
63
65
67
69
71
73
75
77
79
81
83
85
87
89
91
93
95
Source: Chaiyuth 1998
TC in Agriculture
TC in Non-Agriculture