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U.S. Farm Programs and Farmers' Incomes

by

Bruce L. Gardner

Department of Agricultural and Resource Economics
University of Maryland
College Park, Maryland 20742

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The rationales for U.S. farm programs are many, but the chief objective apparent in political debate is to support farm income. It should therefore be a devastating criticism of the programs if they do not raise farm income. Indeed, some farm policy experts believe farm programs do very little for farm income in the long run beyond raising land prices. Yet each farm policy debate in Congress involves estimates that farm income will be significantly affected by the policy adopted, aided and abetted by the testimony of farm policy experts as well as farmers themselves.

This paper attempts to reconcile these apparently conflicting views and to estimate the effect of U.S. farm programs on U.S. farm income over the last 57 years (1933 - 1990).

Background

Here are views from the left, center, and right wings of farm policy expertise:

> Over the long period 1946-70, net farm income measured on a per acre basis for the United States held almost constant, but the value of farm land increased from $66 to $174 per acre. The ultimate gainer from a net income increase in agriculture, whether resulting from an increase in demand, a farm technological advance, or a farm program, is the land owner. Any income gain tends to get capitalized into the limiting input, land, through the competitive process. And that is where the income benefits of the farm programs had to come to rest (Cochrane and Ryan, 1976, p. 371).

"In formulating realistic policies, it is well to recognize that commodity programs do not raise the net income of farm people over the long run". (Tweeten, 1989, p. 419).

"A major degree of agricultural protection contributes little or nothing to the long-run solution of the farm income problem ... high prices alone are meaningless as a long-run farm-income measure;" (Johnson, 1991, p. 215).

The general point of this widely shared view is that income gains "get capitalized into the limiting input, land" (Cochrane and Ryan). The rental value of land increases and this increases the net income of farmers who own land, so the Tweeten quotation overstates the point. Johnson expands the point to say that farm labor returns will also increase to the extent labor is not perfectly elastic in supply to agriculture in the long run.

At the same time, economists who model the effects of one- to five-year farm bills generally agree that higher support levels increase net farm income. In evaluating alternatives for the 1985 farm bill, Galston (1985, pp. 139-40), reported estimates that a move to a free-market policy, as compared to a continuation of the programs in place, would cause U.S. farm commodity prices to fall 9 percent, and this would reduce net farm income by 25 percent ($6 billion) on average in 1986-89. Stanley Johnson et al. (1985) estimated that for 1986-90 a market-oriented policy that caused the prices of basic agricultural commodities to fall by 15-20 percent would cause net farm income to fall 30 percent (pp. 171-2). These are quite large effects.

The most plausible way of reconciling these net income estimates with the earlier quotations is that the numerical estimates, though for a prospective 5-year period, are largely short-run effects that would not persist. But similar results would have been obtained for practically any 5 years during 1933-1990. So are we to say that farm programs generated substantial short-run gains in each of the last 57 years, but little significant long-run gains? This doesn't make sense because all those short-run gains have been received and spent. It's too late to wipe out the 1950s gains in the 1980s.

The proper long-run comparison is between the current situation and what the current situation would be if there had been no farm programs since 1933.

Estimating Income Effects Using U.S. Aggregate Time Series

Consider a restricted profit function for U.S. agriculture in year t:

(1) $$\pi_1 = \pi^*(P_t, w_t, z_t)$$

where P_t, w_t, and z_t are vectors of farm product prices, input price, and fixed factor quantities. The fixed factors include policy variables such as acreage idling requirements or payments to producers.

Our aim is to estimate the effects of the policy-determined elements of z_t on industry profits. Empirical implementation of (1) involves the following:

- We assume that agriculture is a competitive industry with zero expected profits beyond competitive returns to factors of production.
- π_t therefore is the expected value of rents to fixed factors plus random (positive and negative) profits to the farmer as entrepreneur.
- The data which most closely approximate this concept are USDA net farm income plus rents paid to nonfarm landlords. This assumes that no rents accrue to suppliers of purchased inputs, including machinery and structures, or to hired farm workers.[1]

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[1] Use of USDA net farm income data does not preclude machinery and structures being fixed inputs to the farmer in the short run. The reason is that USDA does not subtract actual rental values of machinery or structures from gross revenue to obtain net income. Instead, formula-based depreciation, interest, and repair costs are subtracted. This means short run increases or decreases in rents to suppliers of these factors show up in net farm income (which is where they belong for farm-owned capital).

- The policy variables in z_t typically influence P_t and sometimes w_t (as in the cases of input restrictions or subsidies). Therefore, it is difficult to identify separately the effect of z_t and (non-policy determinants of) P_t.
- In order to identify the z_t effects, exogenous instruments for P_t and w_t are used. The instrumental variables are:

  1) Macroeconomic variables:

     i) rate of growth of real GNP as a business cycle indicator

     ii) rate of inflation (CPI growth)

     iii) a dummy variable for the commodity boom years of World War I, World War II, and 1973-74.

  2) A representative non-farm real wage rate, the weekly earnings of workers in manufacturing and retail trade.

  3) The USDA total factor productivity index.

The representation of agricultural policies requires drastic simplification. Legislation and regulation have established hundreds of policy instruments — market support prices, target prices for determining deficiency payments, acreage set asides, government commodity purchases, export subsidy payments, export credit guarantees, disaster payments, commodity storage payments. And the instruments used and commodities covered have changed over time.

Attempts to simplify the representation of these policies have gone so far as to reduce all of U.S. farm policy to a scalar for each year — the "producer subsidy equivalent" or "aggregate measure of support." These measures aggregate too many disparate items in too arbitrary a fashion. This paper boils the policies down to three primary types of intervention, and ignores all others. Four indicators are used as elements of z_t:

1) A dummy variable = 1 from the introduction of the New Deal, 1933, and = 0 before 1933.
2) Aggregate payments received by farmers from all programs.
3) Acreage idled under set-aside, conservation reserve, or other related programs.
4) The (constant-dollar) value of commodity stocks acquired by the Commodity Credit Corporation, the federal government's price support arm.

The generalized quadratic is used as a flexible function form for estimating equation (1). Suppressing interaction terms on the macroeconomic instruments, the estimated equation is:

$$(2) \qquad Y_t = \alpha + \sum_{i=1}^{5} \beta_i X_{it} + \sum_{j=1}^{4} \gamma_j z_j + \sum_{j=2}^{4} \sum_{k=2, k\leq j}^{4} \gamma_{ij}\, z_i\, z_j + \varepsilon_t$$

where the x_i are the five instruments listed above, the z_1 are the farm policy variables (z_1, the post-1933 dummy, is not included in the z_i z_j variables because the cross products $z_1 z_j$ are collinear with z_j). The results are shown in Table 1.

The hypothesis that the 10 policy variables are jointly not significantly different from zero cannot be rejected at even the 10 percent level: F(10, 64) = 1.1. But this test is hard to pass because the specification loses numerator degrees of freedom needlessly by including so many highly collinear variables. Table 2 shows the predicted effect of programs implied by the γ coefficients of Table 1 in each year since 1933. The estimated effect is:

Table 1. Regression explaining U.S. real net farm income, 1911-1990

equation(2) parameter	description of variable	estimated parameter	"t" ratio
α	Intercept	12513	9.3
β_1	Commodity boom years	8846	7.2
β_2	Inflation rate	12481	1.4
β_3	GNP growth rate	8158	1.2
β_4	Productivity	-72.4	-1.9
β_5	Index of WN	59.0	1.5
γ_1	= 1 after 1932	2162	1.3
γ_2	government payments	-30.3	-0.0
γ_3	acreage diverted	-3.58	-0.4
γ_4	government purchases	-692	-0.8
γ_{23}	interaction term	11	1.4
γ_{24}	"	757	1.0
γ_{34}	"	-4.37	-1.2
γ_{22}	squared term	-1063	-1.0
γ_{33}	"	-.031	-1.5
γ_{44}	"	447	-1.1

$R^2 = .66$

Dependent variable mean: 13,585 (1967 dollars).

(3) $$\hat{G}_t = \sum_i \hat{\gamma}_i z_{it} + \sum_i \sum_{j \le i} \hat{\gamma}_{ij} z_i z_j$$

The mean value of $\hat{G}_t$ is 1,128, meaning that on average the farm programs are estimated to increase net farm income by $1.128 billion (in 1967 dollars), or about 8 percent of mean net farm income.

U.S. net farm income has been almost constant in real terms (Figure 1). In 1910-14 it averaged $13.8 billion and in 1987-90 $12.4 billion. In between real net farm income reached a low of $4.8 billion in 1983 ($5.0 billion in 1932) and a high of $25.8 billion in 1923. The regression trend decline of $16 million (0.1 percent) per year is not significantly different from zero (t = 0.7) and is obviously swamped by the instability of income in Figure 1. The R^2 of .66 in Table 1 is lower than in typical regressions explaining economic time series, and the lack of trend in the dependent variable is probably the main reason. The up-side of this situation is that we have to worry less than usual about spurious correlation of the dependent variable with trending right-hand-side variables.

The lack of a pronounced trend is misleading in that real income per capita was rising, even relative to incomes in the nonfarm sector. Did farm policy play a role in this development? Table 3 shows a regression using the same specification as equation (2) explaining real income per farm (USDA net farm income in 1967 dollars divided by the number of farms as estimated by USDA using the Bureau of the Census farm definition), and also a regression explaining farm household income from all sources as a percentage of nonfarm household income (as published by the Economic Research Service of USDA in Economic Indicators of the Farm Sector).

Table 2. Estimated effects of farm policies on net farm income

Year	$\hat{G}_t$
1933	2031
1934	325
1935	1111
1936	-30
1937	860
1938	945
1939	-1467
1940	169
1941	753
1942	183
1943	612
1944	-224
1945	134
1946	343
1947	2119
1948	2138
1949	2198
1950	3163
1951	4197
1952	2345
1953	2226
1954	2435
1955	2307
1956	1342
1957	2561
1958	1393
1959	692
1960	-211
1961	1
1962	874
1963	-1452
1964	289
1965	1111
1966	3845
1967	1215
1968	287
1969	217
1970	741
1971	930
1972	645
1973	729

1974	2112
1975	1873
1976	1920
1977	1104
1978	1198
1979	1724
1980	1832
1981	1595
1982	1369
1983	-2769
1984	376
1985	615
1986	828
1987	-25
1988	4646
1989	951
1990	571

Figure 1

Table 3. Regressions explaining income per farm, farm as percentage of nonfarm household income, and farm numbers, 1911-1990

equation(2) parameter	dependent variable	farm income per farm		farm/nonfarm income		farm numbers	
α	Intercept	-.69	(2.1)[a]	12.9	(2.4)[a]	9584	(94)[a]
β_1	Commodity boom years	1.79	(6.0)	20.8	(4.3)	307	(3.3)
β_2	Inflation rate	.56	(0.3)	91.9	(2.6)	-2011	(3.0)
β_3	GNP growth rate	2.50	(1.6)	37.7	(1.5)	100	(0.2)
β_4	Productivity	.032	(3.6)	.84	(5.9)	-45	(16.1)
β_5	Index of WN	.020	(2.2)	-.02	(0.1)	-25	(8.6)
γ_1	= 1 after 1932	-.12	(0.3)	-15.4	(2.5)	824	(6.9)
γ_2	government payments	-.20	(0.4)	-9.3	(0.6)	115	(0.8)
γ_3	acreage diverted	-.31	(0.2)	-6.6	(0.2)	-346	(0.6)
γ_4	government purchases	-.19	(0.9)	-2.2	(0.7)	-71	(1.1)
γ_{23}	interaction term	3.30	(1.7)	50.4	(1.6)	555	(0.9)
γ_{24}	"	.22	(1.2)	3.1	(1.0)	86	(1.5)
γ_{34}	"	-1.4	(1.7)	-15.	(1.1)	-494	(1.9)
γ_{22}	squared term	-.25	(1.0)	-3.0	(0.8)	-63	(0.8)
γ_{33}	"	-9.1	(1.9)	115	(1.5)	2131	(1.4)
γ_{44}	"	-.018	(0.2)	.91	(0.7)	-59	(2.0)

[a] t-ratio (absolute value)

R^2 = .84, .85, .99 (left to right)

Dependent variable means: $3.26 thousand (1967), 73.2 percent, 4774 thousand farms.

Figure 2

Note that the non-policy variables show different effects on the per-unit income measure. The non-farm wage rate is highly significant for farm income but is negative in sign, though not significant, for farm/nonfarm income. This is expected because while a rise in nonfarm wage rates indicates improved labor market conditions for farm people, higher nonfarm wages are even better for nonfarm people. The effect of the total factor productivity variable changes from a negative effect on aggregate farm income to a positive effect on per capita farm income, and on farm relative to nonfarm income.

The policy effects are different, too, although this is difficult to see given the varying signs and low significance of the individual policy variables. Aggregating farm policy effects as in equation (3) indicates that on average in 1933-90, policies reduced income per farm by $500.

These results raise the question of how policies influenced farm numbers. A regression explaining farm numbers is shown as the right-hand column of Table 3. The effects of policies shown imply that policies on average in 1933-90 kept the number of farms 723,000 higher than it would have been without the policies. Thus, the finding is that farm policies kept farmers in business and this increased aggregate U.S. farm income but reduced the average farmer's income.

<u>Cross-sectional data</u>

Long-run effects of commodity programs on farm income should be detectable comparing farms which are heavily engaged in producing supported commodities with farms which concentrate more on non-program commodities. Some 47,300 individual records of USDA's Farm Costs and Returns Survey were examined to try to identify farm income effects of commodity programs in 1987, 1988, and 1989.[2] The data records contain revenue and expense information sufficient to calculate the net income of each farm, the commodities produced on the farm, and government payments received.

Because farm programs often raise market prices as well as providing payments, the amount of payments received by farmers is an incomplete indicator of program benefits. The sector-wide indicators used in the aggregate time series include acreage diversion and government purchases, but we do not have farm-specific or commodity-specific measures for cross-sectional purposes. The variable used in the regressions to measure commodity program effects is based on the "producer subsidy equivalent" (PSE) measures developed by the OECD. A commodity's PSE attempts to capture all of the effects of governmental interventions on a commodity's returns — not only direct price support but also input subsidies, transportation subsidies, even the value of research and extension activities. The value of government benefits bestowed on each commodity is then divided by the commodity's market value to obtain a percentage indicator. PSEs thus measured are an inexact measure and have been much criticized. Nonetheless, it is fair to suppose that commodities with large PSEs do well at obtaining government support as compared to commodities with low PSEs.

A support indicator for each farm is obtained by multiplying each commodity's PSE times value of that commodity produced on the farm. Summaring over commodities produced obtains a dollar value of support for the i^{th} farm:

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[2] I am indebted to the Economic Research Service of USDA, and to Gerald Whittaker particularly, for making the FCRS data available and for carrying out the estimation reported in Table 4.

(4) $$G_i = \sum_{j=1}^{M} P_j V_{ij} \quad ; \quad i = 1, N$$

where P_j is the (nationwide) PSE for the j^{th} commodity, V_{ij} is the value of production of this commodity on the i^{th} farm, and G_i is the resulting estimate of support for each of the N = 47,300 farms.

In addition, a dummy variable is added to the regressions, equal to 1 if the farm received government payments (i.e., participated in programs) or zero if no payments were received; as well as an interaction (cross-product) term between participation and the farm's PSE.

Table 4 shows the results of a regression model intended to estimate the effects of commodity support on farm income, holding other economic factors constant. These also differ from the macroeconomic instruments used in the earlier regression, since the macroeconomic variables are the same for all farms. Instead, two farm-specific economic factors are included in the regressions: the gross value of production on the farm, an indicator of farm size; and a measure of the farm's efficiency. The efficiency measure is a cost index, obtained by dividing total cash expenses by the value of production on the farm. Both the size and cost-index variables are statistically significant. The coefficients indicate that as the value of a farm's output rises by $1.00, net cash income rises 19.8 cents; and that a rise in cost per dollar of revenue of 1 percentage point causes net cash income to fall by $11 (or .07 of 1 percent of the $14,450 sample mean of net cash income).

Table 4. Regression Explaining Net Cash Income of 47,300 FCRS farms, 1987-1989

Independent variable	Coefficient	"t" statistic
Intercept	153	1.3
Size of farm	.198	4.3
Cost index	-11.0	2.3
PSE	-.029	0.3
Participation	100	0.3
PSE * Participation	.279	1.6

$R^2 = .11$

Dependent variable mean: $14,450 (nominal).

The effect of farm programs is, at sample mean values: -.029 (9381) + 100 (.275) + .279 (9381) (.275) = $476. However, the $476 effect is not statistically significantly different from zero at the 5-percent significance level. Other specifications of the regressions yielded similar results — positive but small and questionably significant effects of farm programs on farm income. Note that even if the $476 were the correct value for all 2.1 million U.S. farms in 1987-89, the aggregate farm income gains would be just $1.0 billion annually.

Implications of Regression Results

The estimate of a $1.0 billion farm income effect in 1987-89 from the cross-sectional regression is about half the effect of programs is estimated in Table 2 for the 1987-89 average from the time-series regression. This could be explained by farm programs increasing the value of farm assets generally. The idea is that in equilibrium farmers who grow non-subsidized potatoes have to return as much to their resources as farmers who grow heavily subsidized wheat. Otherwise the potatoes growers would switched to wheat. But there are some natural and artificial (policy caused) inpediments to such substitution.

The farmers' gain as estimated in either the cross-sectoral or time series regression is far less than the $15.1 billion average CCC outlays for 1987-89, not to mention the consumers' costs of production and import restrictions. This implies a much larger deadweight loss per dollar transferred to farmer than supply/demand simulation of farm policies have generated.

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