Seeds of Progress?
French Wheat Production, Quality, and Policy

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Abstract

Recent theoretical work suggests that government policies may influence the quality mix of a commodity. This paper provides empirical evidence of quality responses to government policies for wheat policy implemented in France. Analysis is conducted using a detailed data set that includes the class, a measure of varietal importance (area of land used for seed multiplication), quality measures, and experimental yields for each wheat variety grown in France between 1973 and 1999. Results show statistically significant changes in the distribution of wheat produced across quality classes, and in quality and yield indexes, occurring at times of important policy changes.

Keywords: wheat, CAP policy, quality

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Studies typically use a model of a homogeneous good to describe and estimate the price, quantity, and welfare effects of agricultural price policies. However, such a model will fail to account for potentially important policy effects when the commodity of interest is heterogeneous, and the policy distorts the incentives to produce various qualities, and therefore alters the mix of qualities produced under the policy.

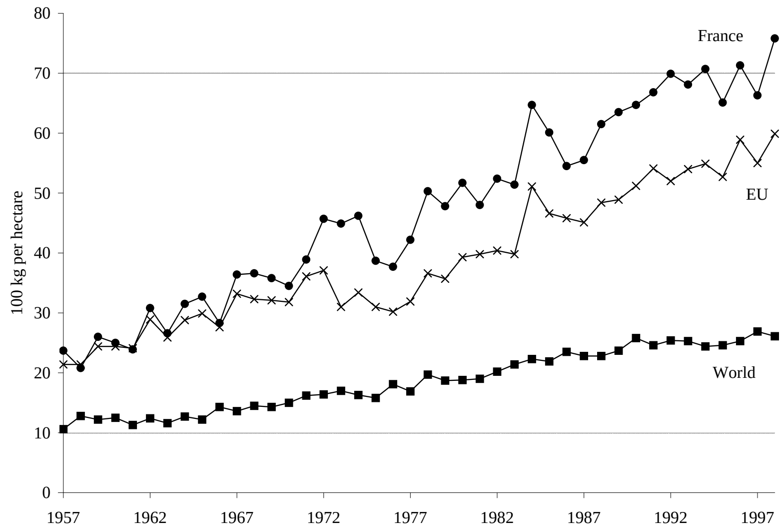
One instance where the quality response to an agricultural policy appears to have been important is the wheat policy implemented in France as part of the Common Agricultural Policy (CAP). The cereals policy component of the CAP has changed over time, but essentially acts as a target-price policy. It is well known that target-price policies create an incentive to increase production, and that the increase in production is amplified over time, with a dynamic response to price. Thus, it is not surprising that wheat production has been increasing in the European Union, in general, and France, in particular (International Grain Council).[1] This quantity response to the policy has exacerbated the costs of the CAP, which have been so large as to induce policy changes designed specifically to reduce its costs (Ingersent and Rayner, 1999).

James (2000) suggested that, in addition to creating an incentive to increase production, policies of this type can also create an incentive for producers to adopt lower-quality, higher-yielding wheat varieties which would further increase the quantity response to the policy, and thus the costs of the program.

As shown in figure 1, wheat yields in France are more than double the world average for most years and have grown faster than world average yields and yields in other major producing countries (figure 2). This paper uses a unique data set to explore the extent to which these yield patterns involved a reduction in quality that can be attributed to the incentives created by the CAP. The analysis shows that statistically significant changes in the distribution of wheat production among classes of wheat as well as structural changes in the time paths of yield and quality indexes occurred at times of policy reform. Further, it is shown these quality changes have substantial implications for the taxpayer cost of the wheat component of the CAP in France.

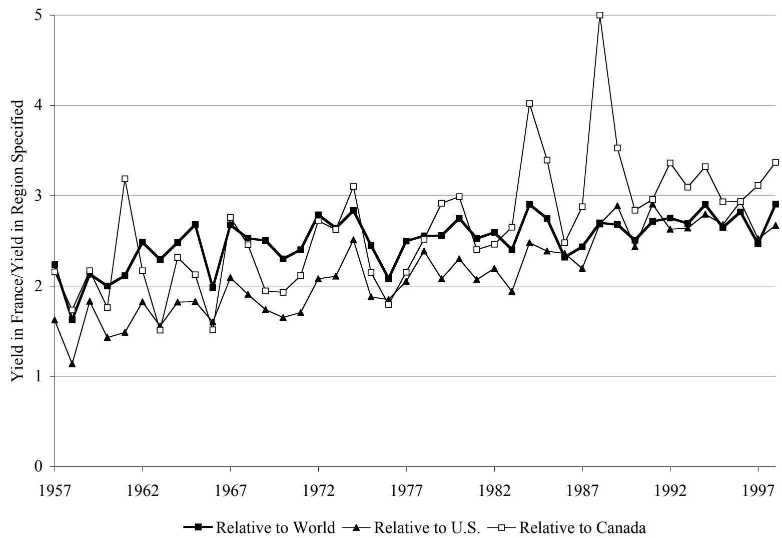
[1]Prior to 1973, the European Community was composed of 6 member countries: Belgium, France, Germany, Italy, Luxembourg, and the Netherlands. In 1973, Denmark, Ireland, and the United Kingdom were added. Greece was added in 1981; Spain and Portugal in 1986; and Austria, Finland, and Sweden in 1995. The European Community became the European Union in November of 1993, when the Maastricht Treaty was enacted.

Figure 1: Wheat Yields in France, the European Union, Major Wheat-Growing Countries, and the World



Source: International Grain Statistics and International Wheat Statistics.
Note: Changes in the composition of the European Union are incorporated in the EU yield data.

Figure 2: Wheat Yields in France Relative to Wheat Yields in the United States, Canada, and the World



Source: International Grain Statistics and International Wheat Statistics.

1 Policy Background

The CAP is a complex policy.[2] The cereals policy of the CAP acts primarily as a set of price supports. Each year, two institutional prices are set, the first of which is the intervention price. The government stands ready to buy wheat at the intervention price, so it acts as a guaranteed producer price. The intervention price is usually above the world price, so export subsidies are offered, so that domestic surpluses may be sold in the world market. The second institutional price is the threshold price, which acts as a minimum import price and has been maintained by the imposition of variable import levies.

In most years, both of the institutional prices bind, and wheat is simultaneously imported and exported in the European Union. Differentiation between the wheat imported and the wheat exported is the only reasonable explanation for this intra-industry trade, as noted by de Gorter and Meilke (1987, 1989) and Meilke and de Gorter (1988).[3] France exports lower-quality wheat, and imports higher-quality wheat (Wilson and Hill 1989, de Gorter and Meilke, 1987 and 1989). Hence, the threshold price acts as the price for high-quality wheat, while the intervention price acts as the price for lower-quality wheat.[4]

The intervention and threshold prices for 1973 through 1998 are plotted in figure 3. Prior to the 1976/77 marketing year, a single intervention price was specified for wheat.[5] Beginning in 1976/77, wheat that was not of bread-making quality qualified for the intervention price of barley (labeled "Intervention Price–Feed Wheat" in figure 3), and a separate reference price was established for wheat of breadmaking quality (labeled "Intervention Price–Bread Wheat").

Between 1976 and 1986, all three institutional prices increased, and production

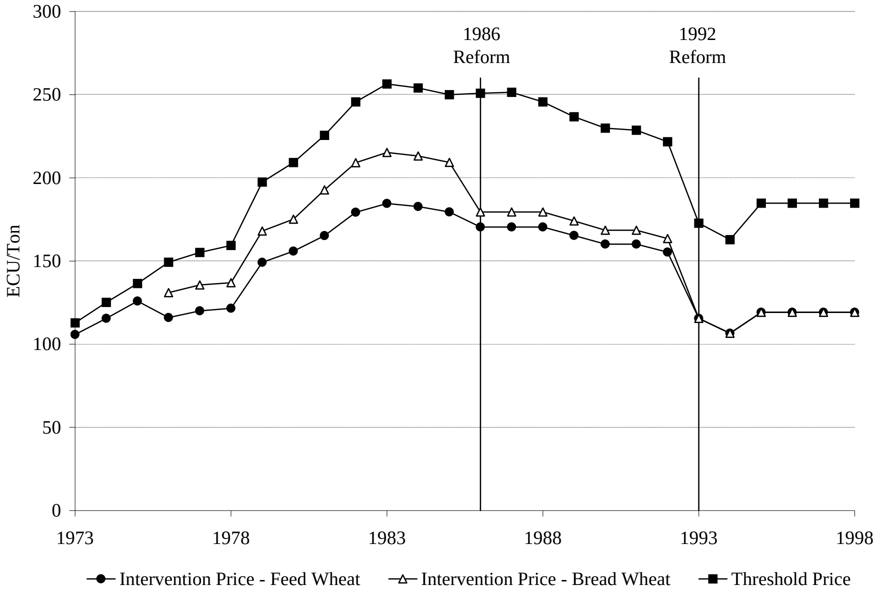
[2]This discussion of the policy is based on de Gorter and Meilke (1987, 1989), Meilke and de Gorter (1988), Wilson and Hill (1989), Swinbank (1997), Fearne (1997), Ingersent and Rayner (1999), and various issues of *The Agricultural Situation in the European Union.*

[3]Border trade may also explain intra-industry trade, but the quantities traded in this instance seem too large to be accounted for by border trade alone.

[4]Wheat attributes that contribute to its breadmaking quality include protein content, gluten content and strength, color, fiber content, cleanliness (i.e., absence of foreign matter), and moisture content.

[5]The marketing year is defined as July to June. The earlier year is used to denote the marketing year in the graphs and tables presented here, i.e., 1976 denotes the 1976/77 marketing year.

Figure 3: Institutional Prices of the Common Agricultural Policy for Common Wheat

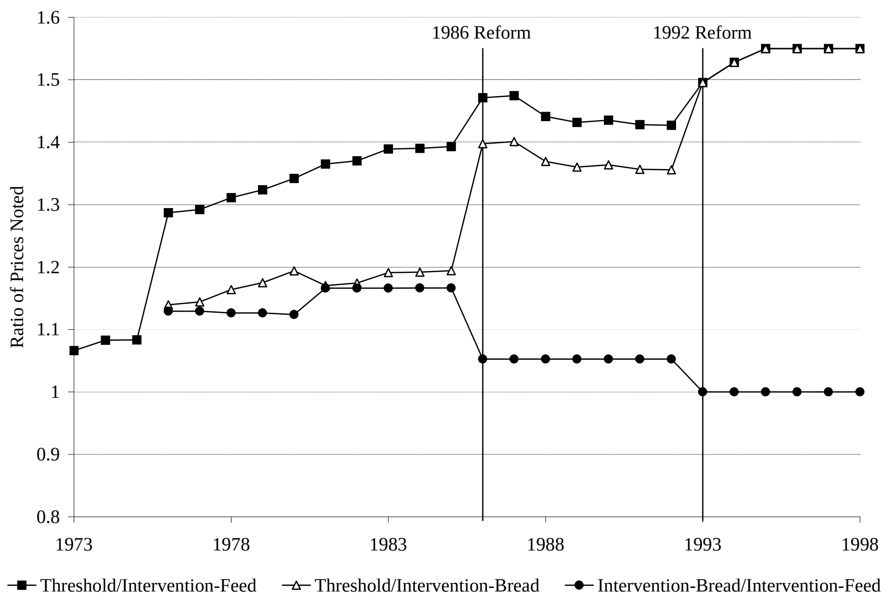


Source: European Commission.

of wheat increased dramatically. By the late 1980s, budget expenditures were a major concern of EU policymakers, and 1986 reforms were designed to curb production (and thus budget costs), and to make producers "feel the realities of the market" (Wilson and Hill, 1989; Ingersent and Rayner, 1999). The intervention prices were reduced in 1986. In addition, a co-responsibility levy was introduced, which taxed producers at a rate that was fixed annually, based on expected production levels. Further policy changes were made in the 1988/89 season, when "stabilizers" were introduced into the co-responsibility levy scheme, increasing the levy and decreasing the following year's intervention price if the aggregate quantity produced exceeded a specified quantity. The effectiveness of these co-responsibility levies has been called into question, however, since agricultural ministers refused to follow through on price cuts (Fearne, 1997). Regardless, the levies were eliminated in 1991.

The relationships among the institutional prices indicate the quality premiums offered in the European Union, as shown in figure 4. Here, two clear patterns emerge. Between the 1976 and 1986 reforms, the threshold price relative to the intervention price for feed wheat grew slowly, and the other two quality premiums were relatively stable. With the 1986 reforms, however, the threshold price relative to the intervention price for

Figure 4: Relative Prices of Low-, Medium-, and High-Quality Wheat under the CAP and in the Rest of the World



Source: European Commission, International Grains Council, International Wheat Council.

bread wheat increased dramatically because of the sharp decrease in that intervention price. For the same reason, the premium for bread wheat over feed wheat decreased sharply. The 1992 CAP reform, first implemented in the 1993/94 season, specified a single intervention price policy for bread and feed wheat, as shown in figure 3. In addition, the price premium defined by the threshold price relative to the single intervention price increased over the first few years of the reform, and has been constant since 1995.

Acreage set-asides have also been part of the CAP reforms (European Commission, 1993). A voluntary acreage set-aside program was introduced in 1988, in which growers who set-aside at least 20 percent of their arable land for 5 years qualified for compensation payments. These voluntary set-aside schemes were in effect until 1997, but mandatory set-asides were also implemented as part of the 1992 CAP reform. The percentage of the set-aside is mandated on a year-to-year basis, and first came into effect during the 1993/94 marketing year (set-aside rates are shown in table 1). In addition to qualifying for the program, growers who set-aside the required percentage of their land received compensation payments based on historical yields in their region.

For each of the policy changes described here, one can imagine the likely effect (if any) it would have on the quality of wheat produced. A unique and extensive data

Table 1: Set-Aside Requirements of the CAP

Marketing Year	Rate of Set-Aside
	percent of arable land
1993/94	15
1994/95	15
1995/96	12
1996/97	10
1997/98	5
1998/99	5
1999/00	10

Source: European Commission.

set allows for a more systematic analysis of the extent to which such quality responses actually occurred. The next section describes the data used for this purpose, and refers back to the institutional prices and policy changes discussed in this section.

2 Analysis of Data for French Wheat Varieties

The data used for this analysis were obtained from various issues of *Semences et Progrès*, a French publication, and include detailed information on the 540 varieties of soft winter wheat registered or planted in France between 1973 and 1999. For each variety, data include the year it was registered, its classification according to three different classification schemes implemented by two different agencies, three measures of quality, and several measures of experimental yields. In addition, the data include the number of hectares used for seed multiplication for each variety and each year (for those varieties that had more than 5 hectares used for seed multiplication) from 1973 through 1999.[6] Each element of the data set is described in turn.

2.1 Wheat Classes

Classification schemes are defined by the Groupe d'Etude et de contrôle des Variétés et des Semences (GEVES) and by the Institut Technique des Céréales et des Fourrages

[6]The share of land area used for seed multiplication for a particular variety is used as a proxy for that variety's share of total land area sown to wheat.

Table 2: Wheat Classification System Used for Analysis

Class	Definition
A	Strong wheat, or improving wheat
B	Wheat of breadmaking quality
C	Suitable for making bread, but lower quality
D	Unsuitable for making bread

Note: Based on classification systems used by GEVES and ITCF.
Source: Various issues of *Semences et Progrès*.

(ITCF).[7] Neither of these classification schemes could be used in this analysis as a single representative scheme because, for a number of varieties, information on classification by only one scheme was available. As a result, the information provided by each agency was used in addition to information regarding the registration status of the varieties to assign each variety to a single class. The resulting classification scheme is described in table 2.

The number of varieties in each class, and two indicators of the prevalence of each class are shown in table 3. As shown in the second column of table 3, 99 of the 540 varieties are classified as strong or improving wheats (i.e., class A), while 132 are class B wheats. The largest category is for class C wheats, which may be used for breadmaking (the wheat will withstand the stress from the machinery used), but are better suited for other purposes. Finally, 33 of the varieties will not withstand the stress from the machinery used, and are classified as not suitable for breadmaking ("impanifiable"). Even though information on the classification variables in the original data set was combined into a single quality class measure, there are still 117 varieties for which no information is available regarding their classification.

By combining this class information for each variety with the seed multiplication data, we can observe changes in the distribution of the area of land used for seed multiplication over the four classes, a measure of the relative importance of each class of wheat. The average number of varieties planted in each class over the period from 1973 to 1999 is shown in the third column of table 3.

[7]GEVES evaluates varieties for purposes of registration, and makes information on characteristics of varieties available to growers. ITCF evaluates varieties, conducts production and market research, and makes the findings available to growers.

Table 3: Details on Number of Varieties and Planting by Class, 1973 to 1999

Class Assigned	Number of Varieties	Average Number of Varieties Planted per Year	Average Percentage of Land Used for Seed Multiplication
	count	*count*	*percent*
A	99	26	27
B	132	30	30
C	159	28	23
D	33	8	6
Unknown	117	20	13

Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

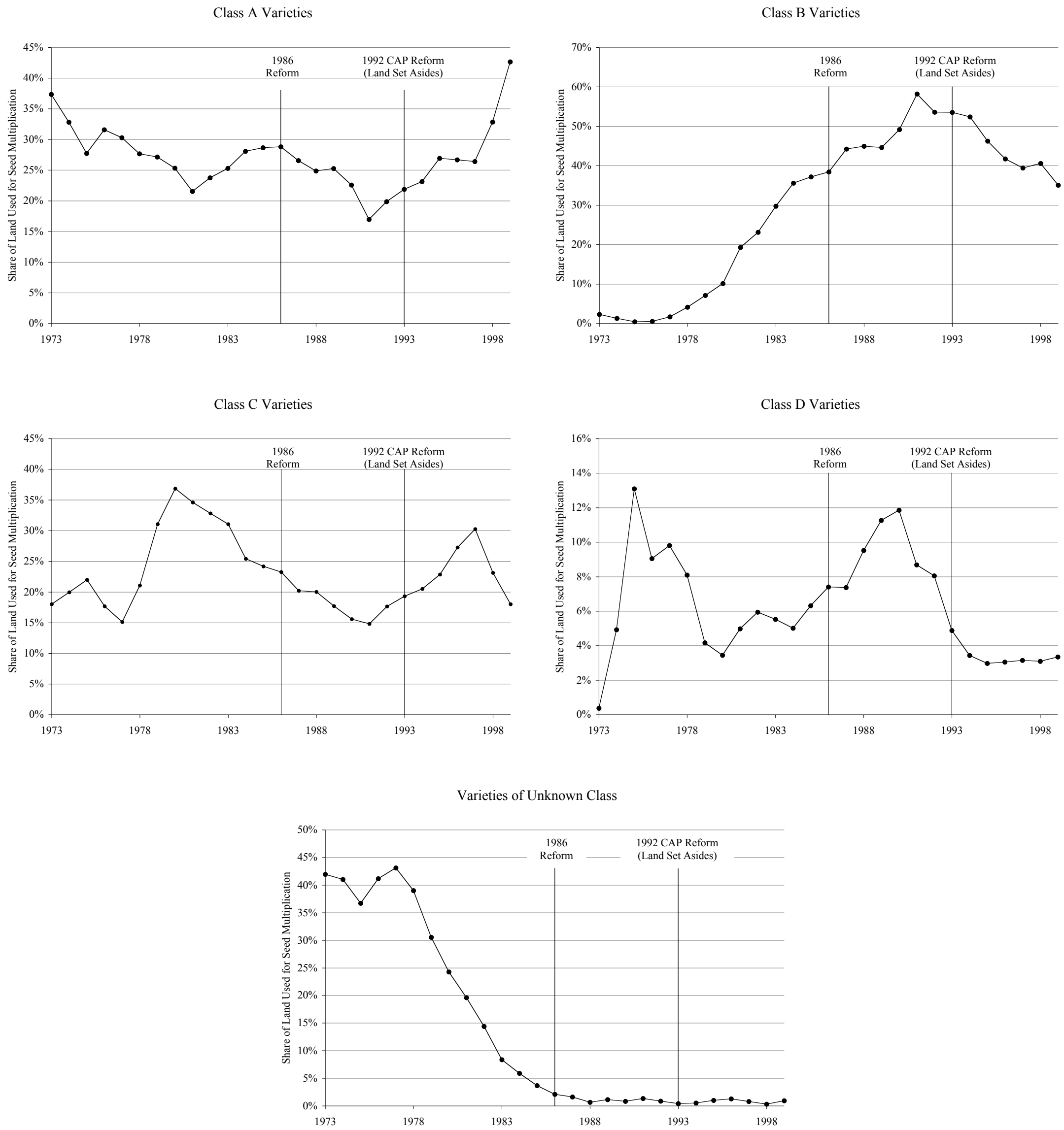
Figure 5 shows the distribution of the area of land used for seed multiplication among wheat classes. The shares shown in figure 5 are calculated including the land used for seed multiplication for varieties of unknown class, and differ somewhat from the shares calculated including only land used for seed multiplication for varieties of known class. These figures reveal changes in the time paths of shares occurring at times of important policy changes.

The first panel of figure 5 shows that the trend in the proportion of land used for class A varieties changed from increasing to decreasing around 1986, and began to increase around the time of the 1992 policy reform. The downward trend between 1986 and 1991 is consistent with the tapering off of the threshold price relative to the intervention price for feed wheat, and the decrease in the intervention price of bread wheat relative to feed wheat occurring over this period of time. The area of land used for each class of wheat as a share of total area of land used for varieties of known class was regressed against intercept and trend dummy variables for the period from 1973 through 1999, 1986 through 1999, 1993 through 1999. The results are presented in table 4.[8]

The results indicate a structural change in the class A share in 1986—while the individual coefficient estimates are statistically insignificant, the joint test that coefficient estimates for slope and intercept dummy variables for the 1986 through 1999 period jointly

[8]The same regressions were run for the share of the number of varieties, and both sets of regressions were run using dummy variables for the period of 1992 through 1999, rather than 1993 to 1999. In addition, results for the same regressions were estimated using shares calculated including varieties and land used for seed multiplication for all varieties (including those of unknown class). The individual estimates changed somewhat, but the signs of the estimated coefficients and the statistical significance of each group of dummy variables changed very little. These results are available in James (2000).

Figure 5: Shares of the Area of Land Used for Seed Multiplication for Each Class of Varieties



Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

Table 4: Regression Results for Policy Effects on Share of Land Used for Seed Multiplication for Varieties in Each Class, Excluding Land Used for Varieties of Unknown Class, 1973-1999

Independent Variable	Dependent Variable: Share of the Land Used for the Class Specified				Tests on System
	A	**B**	**C**	**D**	
Intercept	0.6173	-0.0914	0.3477	0.1264	
	(23.3820)	(-3.7990)	(9.8070)	(5.0940)	
Trend	-0.0299	0.0351	-0.0001	-0.0052	
	(-8.9920)	(11.5990)	(-0.0120)	(-1.6570)	
Dummy Variables for 1986 to 1999:					
Intercept	0.0817	0.0065	-0.1123	0.0241	
	(1.8300)	(0.1590)	(-1.8740)	(0.5760)	
Trend	0.0120	-0.0077	-0.0119	0.0076	
	(1.3190)	(-0.9280)	(-0.9710)	(0.8840)	
Dummy Variables for 1993 to 1999:					
Intercept	-0.0136	0.0034	0.0684	-0.0582	
	(-0.2790)	(0.0770)	(1.0450)	(-1.2710)	
Trend	0.0472	-0.0582	0.0151	-0.0042	
	(3.9380)	(-5.3230)	(0.9390)	(-0.3730)	
R^2	0.8908	0.9632	0.6462	0.3685	
Adjusted R^2	0.8648	0.9545	0.5620	0.2182	
F-Tests			*p-value*		
Coefficients on All Dummies =0	0.0001	0.0001	0.0167	0.3785	0.0001
Coefficients on 86-99 Dummies=0	0.0121	0.6047	0.0225	0.3257	0.0746
Coefficients on 93-99 Dummies=0	0.0030	0.0001	0.3435	0.4027	0.0002

Note: Numbers in parentheses are t-statistics.

equal zero can be rejected at the 5 percent level of significance. The increase in the grow rate after 1992 of class A share reflects the net effect of two opposing forces: a negative effect on quality from the introduction of acreage set-asides, which would suggest a move away from class A varieties to wheat of lower classes[9], and a positive effect after 1993, when the acreage restrictions were relaxed (as shown in table 1); and the quality premium for high-quality wheat (i.e., the threshold price relative to the intervention price) increased. The net result of the two influences was a statistically significant increase in the trend of the proportion of land sown to very high quality, class A wheat.

The trends in the proportion of land used for seed multiplication for class B wheats are opposite those of class A varieties: increasing fairly steadily between the mid 1970s and 1991, and decreasing thereafter. The effects of the 1986 policy reform were statistically insignificant, but those of the 1992 reform were significant. In particular, the rate of growth in the share of land area used for seed multiplication showed a statistically significant decrease after 1992. For class C varieties, there was a statistically significant change in the structure of the time path of the share of land area after 1986, but the change after the 1992 policy reform was not statistically significant. Interestingly, the share of class C varieties seems to move in the same direction as the share of class A varieties, and the share of class D varieties moves in the same direction as the share of class B varieties (as can be seen by combining the relevant coefficients on the trend variables for the particular period of interest). One interpretation of this pattern is that there are actually two levels of grower choices of variety: whether to produce breadmaking or non-breadmaking wheat, and what quality of that type of wheat to produce.

One additional way to measure if any structural change occurred at times of policy reforms is to estimate as a system the same four share equations as those described above. Because the explanatory variables are the same for every equation in the system, the estimated coefficients from the system would not differ from those in table 4. However, estimating the system allows us to test whether all of the estimated coefficients for the dummy variables of interest are jointly equal to zero across all of the equations. The p-value for the null hypothesis that the coefficients for all dummy variables equal zero in all of the equations of the system is presented in the last column of table 4. This p-value

[9]The underlying assumption is that high-quality wheat is more land-intensive, or lower-yielding, than lower-quality wheat. Thus, a land set-aside, which essentially taxes land, taxes the production of high-quality wheat at a higher rate than it taxes production of low-quality wheat.

is 0.0001, indicating that the null hypothesis can be rejected. P-values for similar tests for the coefficients on the dummies for 1986 through 1999 and those for 1993 through 1999 are also reported in table 4. Both p-values are less than 0.10, indicating that we can reject each null hypothesis at the 10 percent level of confidence, and providing further evidence of structural changes. Overall, this analysis provides substantial evidence that wheat price and set-aside policy changes have influenced the relative importance of various qualities of wheat in France.

2.2 Quality Measures

Data on two measures of quality for each variety were also included in the data set.[10] The quality measure used here is the "W" score, which is the wheat's score from Chopin's alveograph test. The test measures dough elasticity, strength, and stretchiness, and indicates the baking quality of the wheat.[11] Usually, "strong" wheat will have a high W score and have 13.5 to 15 percent protein, but some wheats may be rich in protein, yet have only average W scores, and will therefore be of inferior baking quality (Guyonnet 1982).

W scores for the varieties included in this analysis are summarized in table 5. W scores range from 2.5 to 9, with scores in the 6 to 7 range being the most common, both in terms of the number of varieties, and the average share of land area used for seed multiplication. A national index of W scores for each year was constructed as a weighted sum of the individual W scores for the varieties planted in that year, defining weights as the shares of land area used for seed multiplication for varieties that had W scores. This index is shown in figure 6.

The W quality index was regressed on intercept and trend dummy variables, as

[10]Protein scores were available, but only for recently registered varieties, and so they are not used for this analysis. Breadmaking scores were available for 238 varieties. However, because the breadmaking score is more subjective the the W score used here, and fewer varieties have breadmaking scores included in the data set, the inferences drawn from the W scores are probably more accurate, and are the focus of the present analysis. More information on the breadmaking scores is available in James (2000).

[11]Dough is made from the flour, a bubble is formed in the dough, and air is added to the bubble until it explodes. The graph of the pressure in the dough as air is added, and the area under the curve is the W score. The height of the curve indicates the strength of the dough, and is directly related to the capacity of the flour for absorbing water. The length of the curve indicates the stretchiness of the dough. The relationship between the height and length of the curve indicates the balance between the strength and stretch of the dough.

Table 5: W Scores from Chopin's Alveograph Test for French Wheat Varieties

Score	Number of Varieties	Average Percentage of Land Used for Seed Multiplication
	count	*percent*
2.5	5	1.23
3.0	17	5.02
3.5	10	1.62
4.0	22	4.34
4.5	30	10.45
5.0	43	10.44
5.5	41	5.02
6.0	45	16.53
6.5	39	16.32
7.0	35	15.47
7.5	16	5.93
8.0	12	2.92
8.5	4	0.18
9.0	5	1.05
Unknown	216	5.61

Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

was done for the shares of area used for seed multiplication, and the regression results are presented in table 6. The effects of the dummy variables for 1986 to 1999 were statistically significant, as reflected by the p-values for the F-test that both coefficients jointly equal zero. For the period from 1986 through 1999, the time path of the W-score index shifted up and the trend increased relative to the base period, though the latter effect was not statistically significant. Jointly, the coefficients on the two dummy variables were statistically significant, with a p-value of 0.0141 on the joint F-test.

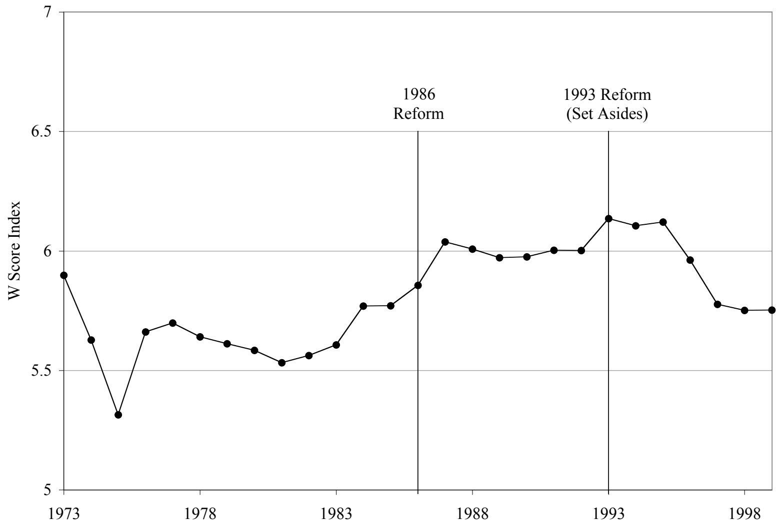
The changes in the time paths for the later time period, from 1993 to 1999, were quite pronounced for the national W-score index. The path of the W-score index shifted up relative to the period of 1986 to 1992, and its slope changed from being positive to negative, with both effects being statistically significantly different from zero and jointly significant. Thus, the level of the W-score index increased, but the growth rate decreased after the 1992 CAP reform. These results provide further evidence of the correlation between policy changes and changes in the quality of wheat produced in France.

Table 6: Regression Results for Policy Effects on W Score Index

Independent Variable	Coefficient Estimate
Intercept	5.6198
	(83.9420)
Trend	0.0024
	(0.2870)
Dummy Variables for 1986 to 1999:	
Intercept	0.2799
	(2.4730)
Trend	0.0095
	(0.4120)
Dummy Variables for 1993 to 1999:	
Intercept	0.2432
	(1.9680)
Trend	-0.0905
	(-2.9760)
R^2	0.7664
Adjusted R^2	0.7108
F-Tests	*p-value*
Coefficients on All Dummies =0	0.0003
Coefficients on 86-99 Dummies=0	0.0141
Coefficients on 93-99 Dummies=0	0.0103

Note: Numbers in parentheses are t-statistics.

Figure 6: National Index of W Scores for French Wheat Varieties



Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

2.3 Yield Data

The data set includes various measurements of experimental yields. Each yield measurement was reported as the percentage of a benchmark variety, for some geographical region and year of experimental trial. In all, there were a total of 181 different measurements (i.e., 181 different combinations of benchmark variety, region, and years of trial), with 5,141 observations. For some of the 540 varieties, several yield measurements for different locations, years, and benchmark varieties were included, while other varieties had few or no yield measurements. Eight different benchmark varieties were used, and they varied over time and geographical region. In addition, some yield measurements referred to fairly specific growing regions of France, while others referred simply to the North or South regions, and others specified no region. Finally, experimental yields were provided by two different agencies (GEVES and ITCF), and some of the reported yields were specific to a certain growing season, while others were not. The number of observations available with each benchmark variety, for each geographical region, and each year of experimental trial are presented in tables 7 and 8 (in addition to coefficient estimates described below).

Each yield observation reflects the influences not only of the variety itself, but also the influences of the benchmark used, the year of the trial (i.e., weather variation), and the growing region. Each observed yield may be represented as Y_{ijtr}, indicating the yield of variety i relative to the yield of a benchmark variety j during time period t in region r. A complete isolation of each effect is not feasible because data are not available for every variety/benchmark/year/region combination. It is necessary to transform the data into a form so that comparisons may be made across varieties, holding constant the influences of the benchmark, weather, and region. A desirable measure of yield would be Y_{iJTR}, which is the yield of variety i relative to a common benchmark J during a common time period T in a common region R. The "common" benchmark, time period, and region, then would be used for all yield measures, so that comparisons among Y_{iJTR} for different values of i, or for different varieties, reflect the differences among varieties, holding all other influences constant.

The ideal transformation of the measures available (the Y_{ijtr} terms) to the desirable measures (the Y_{iJTR} terms) proceeds as follows.[12] Multiplying the observed measures by a time- and region-independent measure of the yield of the common benchmark variety J relative to a time- and region-independent measure of the yield of the observed benchmark variety j would transform the observed measure to one that measures yield relative to the desired benchmark, i.e.,

$$Y_{iJtr} = Y_{ijtr} \frac{Y_J}{Y_j}. \tag{1}$$

Making this transformation for all j benchmark varieties represented in the data (i.e., the 8 varieties listed in the top section of table 7) removes the variation among the observed measures that is attributable solely to variation among the benchmarks used.

The effects of variation in the timing of the trial on the reported yield is removed in a similar manner. Multiplying Y_{iJtr} by a benchmark- and region-independent measure of the yield from the desired base year T relative to the observed year t gives:

$$Y_{iJTr} = Y_{ijtr} \frac{Y_J}{Y_j} \frac{Y_T}{Y_t}. \tag{2}$$

Making this transformation for all years t represented in the data (listed in table 8) removes the variation among the observed yield data that is attributable solely to year-to-year fluctuations. Finally, the variation in the yield data attributable to regional

[12]These derivations are similar to those made by Venner (1997). They omit sampling errors for the sake of notational convenience, but such errors are likely to exist.

variation is removed similarly manner, by multiplying the observed data by a benchmark- and time-independent measure of yields in the desired common region R relative to the same type of measure for the observed region r, i.e.,

$$Y_{iJTR} = Y_{ijtr} \frac{Y_J}{Y_j} \frac{Y_T}{Y_t} \frac{Y_R}{Y_r}. \tag{3}$$

In summary, the reported yield measure, Y_{ijtr}, is transformed into Y_{iJTR} by multiplying the former by benchmark, time, and region yield indexes (the fractions on the right-hand side of equation (3)). These indexes do not exist, but are estimated using the data available. Taking the logarithm of each side of equation (3) gives:

$$lnY_{iJTR} = lnY_{ijtr} + lnY_J - lnY_j + lnY_T - lnY_t + lnY_R - lnY_r. \tag{4}$$

Rearranging terms, the log of the observed yield variable can be expressed as:

$$lnY_{ijtr} = lnY_{iJTR} - lnY_J - lnY_T - lnY_R + lnY_j + lnY_t + lnY_r. \tag{5}$$

The elements on the right-hand side of equation (5) were estimated by regressing observed yields on a set of dummy variables. The estimated regression is:

$$\begin{aligned} lnY_{ijtr} &= \alpha + \sum_{v=1}^{33} \beta_v^V D_v + \sum_{i=1}^{539} \beta_i^I D_i + \sum_{j=1}^{7} \beta_j^J D_j \\ &\quad + \sum_{r=1}^{14} \beta_r^R D_r + \sum_{t=1}^{25} \beta_t^T D_t + e_{ijtr} \end{aligned} \tag{6}$$

where:

$$\begin{aligned} D_v &= 1 \quad \text{for varieties of vintage (year of release) } v \\ &= 0 \quad \text{otherwise} \\ D_i &= 1 \quad \text{for variety } i \\ &= 0 \quad \text{otherwise} \\ D_j &= 1 \quad \text{for yield measurements with benchmark variety } j \\ &= 0 \quad \text{otherwise} \\ D_r &= 1 \quad \text{for yield measurements with region } r \\ &= 0 \quad \text{otherwise} \\ D_t &= 1 \quad \text{for yield measurements for year of trial } t \\ &= 0 \quad \text{otherwise,} \end{aligned}$$

Table 7: Coefficient Estimates for Benchmark and Regional Effects in Yield Regression

Dummy Variable	Number of Observations with Dummy=1	Coefficient Estimate	t Statistic
Intercept	5,141	4.6174	918.29
Benchmark Varieties (8 Varieties, Indexed by j, Base=Soissons)			
Champlein	83	0.1718	8.63
Charles Peguy	27	0.1509	11.30
Etoile de Choisy	42	0.1978	9.56
Fidel	1,196	0.1051	23.01
Gala	473	0.0888	20.71
Soissons	2,187	Base	Base
Talent	525	0.1323	20.45
Thesee	608	0.0586	12.44
Regions (15 Regions, Indexed by r, Base=North)			
North	1,454	Base	Base
South	476	-0.0031	-0.89
East	443	0.0005	0.19
West	91	0.0245	4.97
Paris Basin	445	-0.0124	-4.99
Central	296	-0.0318	-10.86
Central Plains	91	-0.0034	-0.68
Central-East	304	-0.0158	-5.44
Central-West	97	-0.0079	-1.64
Central Humid Zones	67	-0.0048	-0.85
South-East	353	-0.0265	-7.46
South-West	352	-0.0153	-4.34
Poitou-Charentes	294	-0.0020	-0.67
Bretagne and Pays de Loire	343	-0.0060	-2.12
No Region Specified	60	0.0148	1.62

where “vintage” refers to the year of release of the variety or the earliest year it was grown.

The coefficient estimates and summary statistics for the regression are shown in tables 7 through 10. The last three sets of dummies on the right-hand side of equation (6) measure the effects of the different benchmarks, years, and regions represented in the data. The estimates for each of the benchmark and region effects are included in table 7. The base, or common, benchmark variety is Soissons, since it had been the benchmark variety for many of the yield observations (2,187 observations reported yields relative to Soissons, as shown in the second column of table 7). Not surprisingly, the coefficients estimated for all of the benchmark varieties were statistically significant and positive. Soissons has been a major variety for the last ten years, and the positive coefficients for the other benchmark varieties suggest that it has out-performed many of the other varieties. If a yield observation reported relative to some other benchmark variety is larger than one reported relative to Soissons, then that other benchmark variety is inferior to Soissons. The joint test that all of the coefficients on the dummy variables for the benchmark varieties were equal to zero yielded an F-statistic of 93.63 (shown in table 10), so that the null hypothesis that the benchmark dummies have no explanatory power can be rejected.

Estimates of regional effects are included in the second section of table 7. Of the 14 coefficients estimated, 7 of them were statistically significantly different from zero at the 5 percent level. Most of the signs on the coefficients are negative, suggesting that, *ceteris paribus*, yields tend to be lower in the northern region (the base region) than in other regions. The F-statistic for the null hypothesis that the coefficients for all regional dummies equal zero is 16.82, as shown in table 10.

Estimates of the coefficients on the dummy variables for the timing of the trials are shown in table 8. The two agencies reported the timing of the experimental trials differently. Yield data provided by ITCF specified the season of the experimental trial from 1984/85 through 1996/97. The information for the dummies corresponding to those time periods are presented first in table 8.[13] Yield data provided by GEVES did not state specifically the year of the trial, so a separate set of dummy variables was used for

[13]However, the data provided by ITCF for 1997 and 1998 were not explicit about the growing season, though it seems fairly clear that 1997 corresponds to the 1997/98 growing season, and that 1998 corresponds to the 1998/99 growing season.

Table 8: Coefficient Estimates for Year-of-Trial Effects in Yield Regression

Dummy Variable	Number of Observations with Dummy=1	Coefficient Estimate	t Statistic
Year of Trial (26 different categories, Indexed by t, Base=None Specified)			
From ITCF:			
1984-5	286	-0.0121	-2.36
1985-6	257	-0.0431	-8.45
1986-7	358	-0.0078	-1.73
1987-8	384	-0.0033	-0.74
1988-9	347	-0.0466	-10.13
1989-0	327	-0.0703	-15.05
1990-1	251	-0.0178	-4.01
1991-2	228	-0.0049	-1.08
1992-3	246	-0.0042	-0.91
1993-4	223	0.0030	0.64
1994-5	257	0.0008	0.17
1995-6	260	0.0221	4.64
1996-7	306	0.0350	7.44
1997	31	0.0383	4.10
1998	249	-0.0267	-5.14
From GEVES:			
1977	16	-0.0092	-0.39
1978	19	-0.0015	-0.06
1979	20	0.0035	0.17
1980	80	0.0033	0.17
1981	79	-0.0246	-3.00
1982	90	-0.0164	-2.07
1983	110	-0.0184	-2.45
1984	110	-0.0094	-1.55
1984(2)	97	-0.0178	-2.34
1985	97	-0.0194	-3.59
None Specified	416	Base	Base

the year of publication.[14] Of the 25 coefficients estimated, 14 of them were significantly different from zero, and the null hypothesis that they were all jointly equal to zero can be rejected, based on information presented in table 10.

The first set of dummy variables used to draw inferences about varietal variation is constructed according to the vintage of the variety.[15] Some varieties were quite old, having been registered or planted as early as 1950; others were registered as recently as 1998. An individual dummy variable was used for each vintage, as listed in table 9. The coefficients on the vintage dummies are expected to reflect the improvement of varieties over time. This can be seen in table 9, where the dummies for vintages earlier than 1988, the base value for the vintage effect, have negative coefficients (reflecting lower experimental yields), and all but one of the dummies for varieties of a more recent vintage have positive estimated coefficients. Further, all but 2 of the 33 estimated coefficients are significantly different from zero (one of them being the negative coefficient on the 1989 dummy), and the hypothesis that all coefficients are jointly equal to zero can be rejected. Thus, there is persuasive evidence that the year of release, or vintage, is a major factor explaining variation in the yields of different varieties, and that varieties have improved in terms of yield over time.

The final set of dummies is included to capture the pure varietal effects—dummies for individual varieties. The inclusion of both vintage dummy variables and variety-specific dummy variables created an exact linear relationship between each vintage dummy variable and the varietal dummies for varieties in that vintage. To address this perfect multicollinearity, for each vintage, the variety with the most yield observations was chosen as the base variety, and no variety-specific dummy was included for it. As a result, the coefficients presented in table 9 actually include some variety-specific effects as well as the pure vintage effect. The coefficients for the variety-specific dummy variables are not included here because 324 were estimated. Of those, 147 were significantly different from zero. Further, the null hypothesis that all of the coefficients on the varietal dummies are zero can be rejected.

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[14]Yield data provided by GEVES after 1985 did not change over time (i.e., the same yield index was published for several years). These are treated as if no date were specified, which is the base case for the time dummies.

[15]Similar effects have been estimated by Godden and Brennan (1994) and Venner (1997).

Table 9: Coefficient Estimates for Year-of-Release (or Vintage) Effects in Yield Regression

Dummy Variable	Number of Observations with Dummy=1	Coefficient Estimate	t Statistic
Vintage (34 Vintages, Base=1988)			
1950	9	-0.2292	-14.95
1952	1	-0.3266	-7.67
1955	8	-0.1938	-12.07
1962	2	-0.2678	-8.76
1963	8	-0.2978	-18.04
1964	19	-0.1924	-17.38
1966	1	-0.1774	-4.17
1969	20	-0.1760	-13.57
1970	10	-0.1501	-10.46
1973	69	-0.1267	-17.95
1974	50	-0.1655	-18.33
1975	8	-0.2047	-9.44
1976	118	-0.0843	-10.88
1977	170	-0.0846	-12.60
1978	226	-0.1059	-15.41
1980	216	-0.1101	-17.41
1981	38	-0.0918	-7.35
1982	294	-0.0779	-11.98
1983	340	-0.0152	-2.62
1984	252	-0.0731	-8.95
1985	402	-0.0268	-3.61
1986	361	-0.0445	-7.15
1987	153	-0.0586	-7.41
1988	360	Base	Base
1989	284	-0.0144	-1.90
1990	181	0.0134	2.13
1991	273	0.0218	2.79
1992	252	0.0759	11.40
1993	159	0.0511	5.72
1994	303	0.0526	6.24
1995	261	0.0274	2.86
1996	3	0.0988	4.01
1997	165	0.0163	1.57
1998	125	0.0471	3.47

Table 10: Summary Statistics for Yield Regression

R^2		0.6522	
Adjusted R^2		0.6225	
F-Test for Coefficients on:	F-Statistic	Num. DF (Den. DF=4,736)	P-Value
Benchmark Dummies Equal Zero	93.63	7	0.0001
Regional Dummies Equal Zero	16.82	14	0.0001
Year-of-Trial Dummies Equal Zero	26.59	25	0.0001
Vintage Dummies Equal Zero	59.53	33	0.0001
Varietal Dummies Equal Zero	10.52	324	0.0001

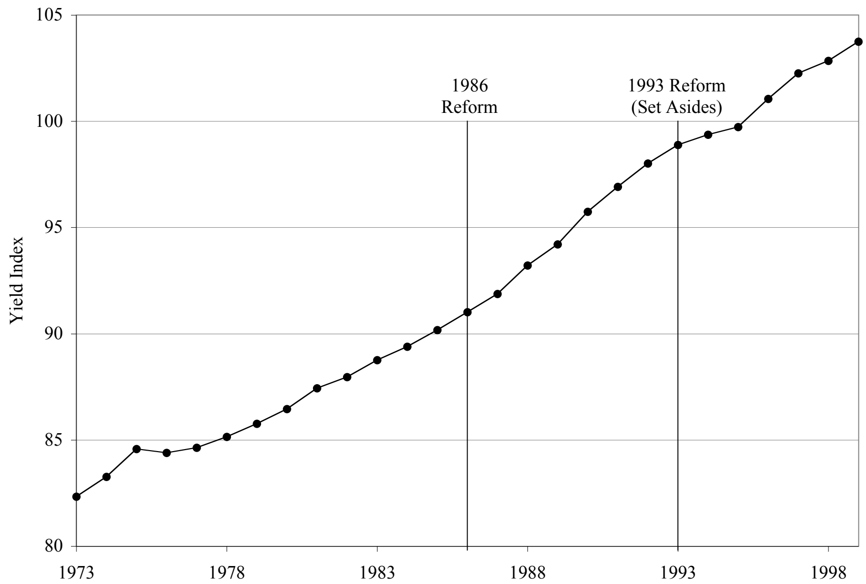
These regression results permit the construction of individual yield measures that include the variety-specific information relative to a common or base numeraire benchmark variety, region, and time of the trial. This is done by calculating the fitted values from the regression, setting all of the benchmark, region, and time dummies equal to zero, while keeping the vintage and variety-specific dummies at their original values. The resulting measure, referred to as the fitted yield, is comparable across various varieties. A national index of these fitted yields was calculated in the same manner as the W score and panification indexes discussed in section 2.2 (as a weighted sum of the variety-specific fitted yields, using shares of land used for seed multiplication as the weights). This index is intended to reflect the change in average wheat yields in France attributable to varietal improvement effects only, weighted by the commercial importance of individual varieties, and is plotted over time in figure 7.

Table 11 reports coefficient estimates for a regression of the yield index over time, similar to those described above. These results show that there were statistically significant changes in the path of the yield index occurring at times of policy changes. In particular, the growth rate of the yield index decreased, relative to the 1986 to 1999 time period, after the implementation of acreage set-asides, and the increases in quality premiums shown in figure 4.

Table 11: Regression Results for Policy Effects on Yield Index

Independent Variable	Dependent Variable: Yield Index
Intercept	81.8690
	(439.8390)
Trend	0.6152
	(26.2320)
Dummy Variables for 1986 to 1999:	
Intercept	-0.2344
	(-0.7450)
Trend	0.5839
	(9.0910)
Dummy Variables for 1993 to 1999:	
Intercept	-0.3347
	(-0.9740)
Trend	-0.3387
	(-4.0060)
R^2	0.9982
Adj'd R^2	0.9978
F-Tests	*p-value*
Coefficients on All Dummies =0	0.0001
Coefficients on 86-99 Dummies=0	0.0001
Coefficients on 93-99 Dummies=0	0.0014

Figure 7: Yield Index for French Wheat Varieties

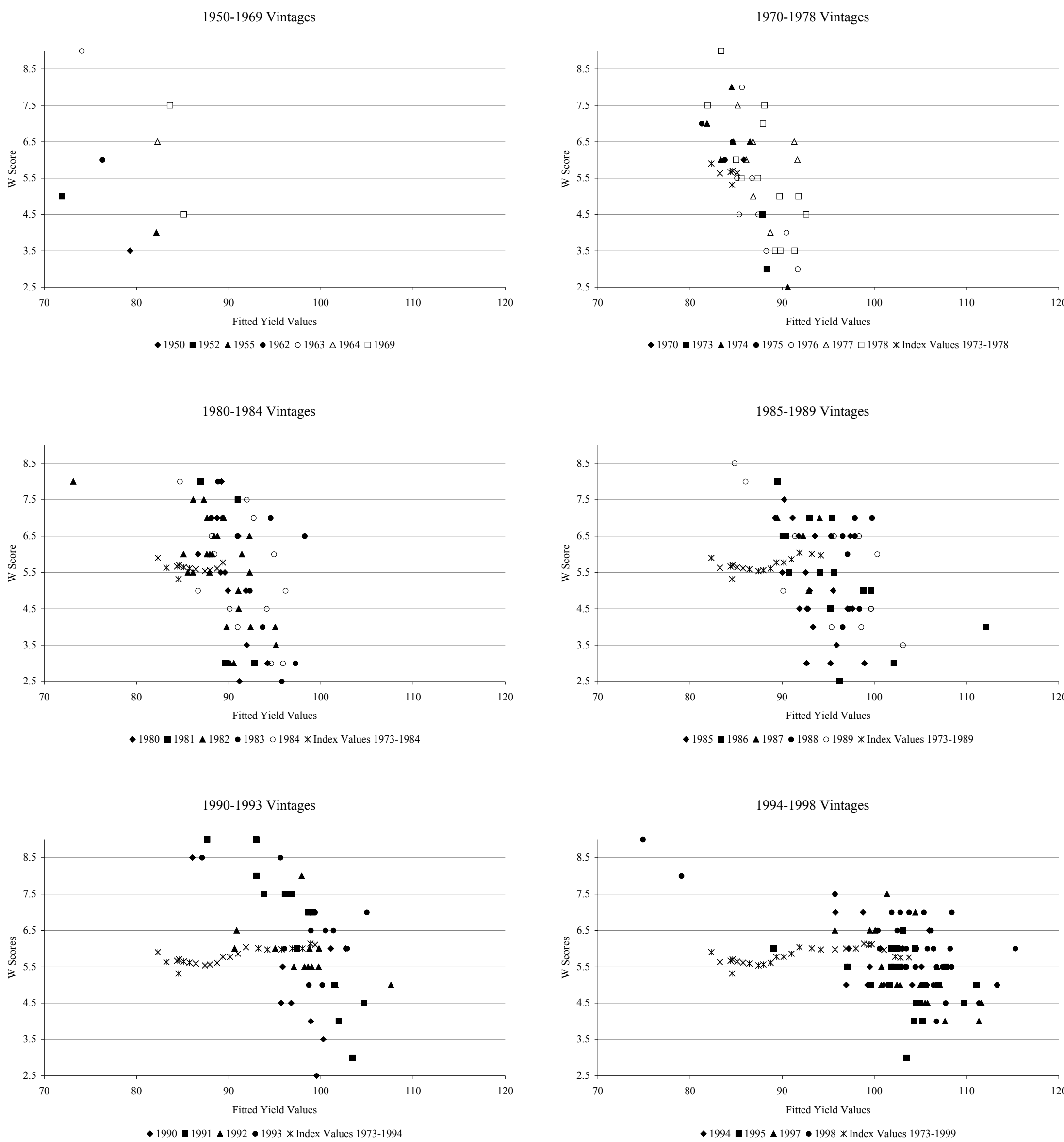


Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

2.4 Linking Yields and Quality

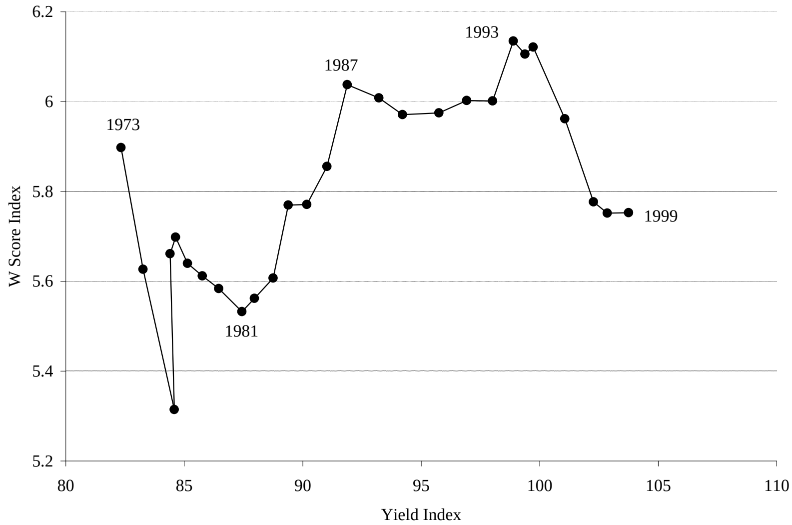
Combining the yield and the quality information described in the previous two sections, we can investigate the tradeoff between quality and yield. This tradeoff is demonstrated in figures 8 and **??**. Figure 8 plots the fitted variety-specific yields against the corresponding W score, for those varieties for which both measures were available. The points farthest from the origin represent a production possibilities frontier, specifically a W score-yield possibilities frontier. The estimated coefficients of the vintage dummies from the previous section suggest that this frontier has shifted out over time—at least in the yield direction. For this reason, the varieties are sorted according to their vintage, and data for groups of vintages are plotted in separate panels. Data from the earliest vintages are plotted in the first panel of figure 8, and the subsequent panels contain the data for successively more recent vintages. In addition, the W-score and yield indexes are plotted for the years up to the most recent vintage in each panel. It is clear that, on average, the W-yield frontier is shifting out and perhaps up over time, implying that newer varieties are improving in at least one dimension. The values for the national yield index and W indexes also increase over time, as indicated by the "X"s in figure 8. But it can be seen that the improvements have been biased in favor of yield rather than W-scores.

Figure 8: W Scores and Yields for French Wheat Varieties, by Vintage



Source: Author's calculations, based on data from various issues of *Semences et Progrès*.

Figure 9: W Score Index Plotted Against Yield Index over Time



Source: Author's calculations, based on data from various issues of *Semences et Progrès.*

The relationship between the W-score index and the yield index is further illustrated in the upper panel of figure 9. In this figure, values for the W-score index are plotted against the yield index, as in figure 8, but the variety-specific observations are omitted so that the path of the W-yield tradeoff may be more easily observed. This figure further illustrates the changes in the overall path of the national quality and yield indexes over time, with the W-score index increasing rather sharply relative to yields prior to 1987, and decreasing after 1993.

3 Budgetary Implications of Quality Changes

To investigate the implications of the observed quality changes for the estimated taxpayer costs of the policy, a measure of taxpayer costs is constructed and its true value is evaluated (using observed data), and then re-evaluated under various counterfactual assumptions. Each set of assumptions is designed to represent a different kind of constant-quality assumption that might be used to formulate a prediction of the next year's taxpayer cost of the policy. Thus, for each year between 1973 and 1998, the taxpayer cost measure for the following year is estimated based on information available, according to the various counterfactual scenarios discussed in more detail below.

The measure used for taxpayer costs of the policy is specified on a per-hectare basis in order to abstract from acreage responses to the policy. For each class, the import levy collected or the export subsidy paid is calculated as the difference between the EU and ROW prices for that class of wheat. The taxpayer cost per hectare of each class of wheat is equal to the export subsidy per ton or the negative of the import levy, depending on the class, multiplied by the class-specific yield index (thus, it is a normalized measure, since the national yield index is normalized to 100). The net taxpayer cost or benefit per hectare is a weighted sum of the class-specific measures, using as weights the class-specific shares of the area of land used for seed multiplication.[16] For the class-specific prices, the prices of high-quality wheat discussed in section 1 are used for class A production, those for medium-quality wheat are used for class B production, and those for low-quality wheat are used for wheat in classes C and D. [17]

As indicated in the previous sections, there are two dimensions of changes in wheat quality occurring over time: the change in the distribution of production across wheat classes, and the change in the quality of wheat in each specific class. Accordingly, the effects of each of two types of errors are explored: the error from making incorrect assumptions about the distribution of land area among classes of wheat, and the error from making incorrect assumptions about yields. Errors are determined by specifying assumptions upon which the estimated shares of production or yields might be based if one were to ignore quality changes, calculating the taxpayer cost measure based on those assumptions, and comparing it with the "true" taxpayer cost measure which is calculated using the actual shares of land and the actual class-specific yield indexes. Alternative assumptions about the shares of production among classes and the yield index for each class are summarized in table 12.

In order to determine the effects of making incorrect assumptions about the quality of each class of wheat, the taxpayer cost measure is evaluated using the observed class-specific shares of land area assuming that the yield indexes do not change from the previous year (scenario 2 in table 12). However, this may over-simplify assumptions made by analysts and policymakers, since even if they ignore quality changes they are

[16]Recall that the share of the area of land used for seed multiplication for each class is used as a proxy for the share of land sown to that class of wheat. These terms are used interchangeably here, with the caveat that they may differ.

[17]World prices were defined as equivalent import or export prices at the EU border. The export price for low-quality wheat is the f.o.b. price of U.S. soft red wheat at the Gulf port. The export price for medium-quality wheat is the f.o.b. price of U.S. No. 2 winter wheat, and the import price for high-quality wheat is the c.i.f. price of U.S. dark northern spring wheat, 14% protein, in Rotterdam.

Table 12: Alternative Scenarios for Evaluating Taxpayer Cost Measure, and Corresponding Percentage Errors in the Estimated Annual Budget Costs of the Wheat CAP in France

	Measures Used for:		Annual Percentage Errors	
Scenario Number	Share of Land Area By Class	Yield Index By Class	Average	Standard Deviation
1	Correct	Correct	0.00	0.00
2	Correct	Same as Previous Year	0.26	6.85
3	Correct	5-Year Moving Average	5.19	19.74
4	Same as Previous Year	Correct	28.50	188.49
5	5-Year Moving Average	Correct	119.34	345.94
6	Same as Previous Year	Same as Previous Year	28.52	192.89
7	5-Year Moving Average	5-Year Moving Average	121.11	354.27

likely to be aware of changes in yields occurring over time. In order to incorporate information regarding yields of recent years, the class-specific national yield indexes are assumed to be equal to the average over the previous 5 years in scenario 3. To determine the error from making incorrect assumptions about the distribution of land area sown to the different classes of wheat, the taxpayer cost measure is evaluated using the observed class-specific national yield indexes, assuming that no change in the distribution of land area occurred from the previous year (scenario 4), and using a 5-year moving average for each share (scenario 5). Finally, the assumptions about the shares of land area and the yield index are combined in scenarios 6 and 7. Scenario 6 in table 12 uses the previous year's values for both the shares of land sown to each class and the yield index in each class, and scenario 7 uses 5-year moving averages for both measures.

The differences between the "true" taxpayer cost measure per hectare, and the estimated taxpayer cost measure per hectare under the different scenarios are summarized in table 12. Using TS_{true} to represent the true taxpayer surplus and TS_i to denote the estimated taxpayer surplus under scenario i, the percentage error in the taxpayer cost measure of the policy is calculated as:

$$\% \text{ Error} \quad = \quad \frac{TS_i - TS_{true}}{|TS_{true}|}. \tag{7}$$

A positive percentage error can indicate an over-estimate of taxpayer benefits (when $TS_{true} > 0$) or an under-estimate of taxpayer costs (when $TS_{true} < 0$). A negative percent error can indicate an under-estimate of taxpayer benefits (when $TS_{true} > 0$) or an over-estimate of taxpayer costs (when $TS_{true} < 0$). Of the 26 years for which the true taxpayer cost per hectare was calculated, it was negative in 17 years and positive in 9 years.

A few general observations can be made about the percentage errors in the estimated taxpayer cost measure, based on the summary information presented in table 12. First, the percentage errors are quite large. For four scenarios, the average of the annual percentage error is well above 25 percent over the time period considered. Indeed, the average annual percentage error is over 100 percent for two scenarios. Some of the larger percentage errors in table 12 reflect a large error combined with a small base value, $|TS_{true}|$, and may be a bit misleading.[18] Second, the errors are much larger when incorrect assumptions are made about the distribution of land area, compared with the errors from making incorrect assumptions about the values of the yield indexes. This suggests that the first dimension of changes in quality—shifts in the distribution of production—is more important than the second—changes in the quality of specific classes—in determining the errors in the estimated taxpayer costs of the policy caused by ignoring quality responses.

4 Concluding Remarks

This analysis suggests that quality responses to agricultural policies do occur. Several changes in the time paths of the relative importance of each of four general classes of wheat, a quality index based on dough quality, and a yield index are found to correspond with important changes made to the cereals policy of the CAP. Further, when these quality changes were ignored, errors in the estimated policy effects were be quite large, on average. This was demonstrated with errors in estimated taxpayer costs of the wheat component of the CAP in France, calculated under various assumptions regarding

[18]The "true" value of the taxpayer cost measure, and the estimated values under each scenario, are included in James (2000), as are tables containing each error, expressed as a percentage of net taxpayer cost (used for the calculations in table 12), of gross taxpayer cost (without incorporating the income from import levies), and of the total value of wheat production (valued at EU prices).

quality. Even under very reasonable assumptions, these proportional errors were quite large, averaging in excess of 25 percent of the true taxpayer cost.

These errors correspond to incremental changes in yields and in the distribution of production across qualities, given an existing policy. We would expect the implied errors from ignoring quality to be larger when comparing a policy outcome to a counterfactual free-market outcome, so that changes in the policy occur in addition to changes in quality and yields. Moreover, over time, we would expect these errors to be bigger as yield and the distribution of production among qualities change in response to the policy.

Several complications arise if the goal of an analysis is to compare policy and no-policy outcomes. First, in order to estimate the quality effects of a policy, a more explicit link must be made between quality and policy variables. While this is done in the analytical models (James 2000), one's ability to implement this empirically is limited by the nature of the data available. For instance, while many of the relationships noted in previous sections give us an idea about the substitution possibilities among various qualities of wheat in the production process, it is not clear how to use the data to estimate the elasticity of transformation between classes of wheat. Other challenges include incorporating changes in the world market prices of the various qualities, and incorporating acreage responses to the policy. It seems quite possible that incorporating both of these influences would tend to make the errors from ignoring quality in such a setting even larger than the measure used here indicates.

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