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CEREAL PRICES IN THE EUROPEAN COMMUNITY:

A POLICY TRANSMISSION MODEL

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ABSTRACT

"Cereal Prices in The European Community (EC): A Policy Transmission Model"

EC cereal policy allows cereal market prices to fluctuate between intervention and threshold prices, according to local supply/demand conditions. To capture the impact of this policy on EC cereal prices, a varying-parameter transmission model is specified and estimated for market wheat prices in France, the Federal Republic of Germany and Britain.

Cereal Prices in The European Community: A Policy Transmission Model

In each country of the European Community (EC), consumers, food processors and livestock producers purchase cereals at market prices that fluctuate between the intervention (floor) and threshold (ceiling) prices (see Figures 1 and 2). Since 1976,[1] by widening the gap between ceiling and floor prices, the EC political authorities allowed regional market prices to be more responsive to local supply/demand conditions (Harris et al.). Thus, an oversupply of cereals characterized by active intervention buying implies that prices are close to the intervention level. On the other hand, in grain deficit regions, an excess demand partly filled by either intra-EC trade or imports from third countries implies that the wholesale price is close to the threshold price.

The objective of this paper is to present a simple model framework which, although linking market prices for cereals to policy-determined prices, will reflect this reaction of market prices to regional market conditions within the EC. For this purpose, a logistic varying parameter model is developed and applied to annual and monthly prices of wheat in the three major wheat producing countries of the EC: France, West Germany and the United Kingdom.

Econometric results presented in Section IV suggest that the evidence of imperfect transmission of policy prices obtained by Colman with a slightly different model for the United Kingdom varies among EC countries. This finding is also dependent upon the choice of the data sample and model specification adopted to represent the transmission of policy cereal prices. In the final

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1. Due to the rigidities of the initial EC cereal price policy regime which, in practice only, allowed regional prices to differ by transportation costs, it was gradually phased out between 1974 and 1976 in favour of a more flexible price system which allowed a response to local supply/demand conditions. At the same time, The EC Council of Ministers improved their cereal polciy through two fundamental policy changes (institution of the SILO system). They introduced, firstly, common intervention prices for feed grains, and secondly, a reference price for bread-making wheat, now set at a higher level.

section, the main econometric findings are summarized and implications of adopting this general price transmission model to evaluate EC cereal policies are discussed.

Conceptual Model:

To link market cereal prices to the corresponding policy prices and regional (or local) supply/demand market conditions, the following linear relationship for a given EC country and period, t, can be used:

(1) $$WP_t = \theta + \beta[\alpha_t PIN_t + (1 - \alpha_t) PTH_t]$$

where: WP_t, PIN_t and PTH_t are the market, intervention and threshold prices expressed in local currency, respectively; θ and β are estimable parameters which are assumed to be positive; and α_t is a varying parameter which takes values between [0 1].

Equation (1) links the wheat market price to a `blend policy' price made up of a combination of the intervention and threshold prices, weighted respectively by coefficients α_t and $(1 - \alpha_t)$, which in turn respond over time to local supply/demand market conditions. Hence, when local supply is abundant, α_t tends to one and the blend policy price tends to PIN_t, implying that the market price is most responsive to variation in the intervention price. On the other end of the spectrum, extreme deficit regions would be typified by a value of α_t close to 0, leading to a direct response of cereal market prices to threshold prices.

The response pattern attributable to the coefficient, α_t, can be formalized by relating the former to a proxy variable, X_t, representing local supply/demand conditions. Such a relationship is written as:

(2) $$\alpha_t = k(X_t)$$

with $\alpha_t \dashrightarrow 1$, when X_t is very large; and $\alpha_t \dashrightarrow 0$, when X_t is small or negative. Furthermore, the function $k(X_t.)$ is non-linear, monotonic and asymptotic to 0 and 1. The properties of this function k(.) imply that its first-order partial derivative with respect to X_t is positive and tends to zero when X_t approaches infinity.

Replacing α_t by (2) in (1) gives an estimable non-linear and varying-parameter relationship between market and threshold and intervention prices which, after rearrangement, can be expressed as:

(3) $$WP_t = \theta + \beta[PTH_t + k(X_t)\ (PINT_t - PTH_t)]$$

This equation, which is the fundamental conceptual model adopted to explain the linkage between market and policy cereal prices, possesses a certain number of interesting features that are worth discussing.

First, differentiating expression (3) with respect to X_t yields a specific relationship and inverse relationship between the market price for wheat and the latter explanatory variable. In fact, as $\partial k/\partial X_t > 0$, and $[PIN_t - PTH_t] < 0$, the following inequality holds:

(4) $$\frac{\partial WP_t}{\partial X_t} = \beta\ [PIN_t - PTH_t]\ \frac{\partial k}{\partial X_t} < 0$$

indicating that the local market price of wheat is a decreasing (increasing) function of the surplus (deficit) position in wheat of the EC region or country under consideration.

Secondly, it is expected that the transmission paramater β will be close to one. However, as pointed out by Colman (1985), the transmission between policy and market prices may not be one-to-one. Imperfect market structures, increasing marketing margins, the timing of marketing operations and institutional rigidities specific to each EC country could cause an imperfect transmission between market and policy prices. Consequently, it is hypothesized that the response coefficient β will take values smaller than one.

Also, while expression (3) is well-suited for EC countries such as the United Kingdom, France and the Federal Republic of Germany, where the total demand is largely filled by domestic production, it may not perform well for small countries like Belgium and the Netherlands, which are in a permanent deficit situation. For the latter cases, we can assume that local supply/ demand conditions do not change much over time, and that the function k(.) can be reduced to a positive scalar which is smaller than one. As a result, the policy transmission model becomes a constant price linkage specification and is similar in spirit and scope to the model estimated by Colman and Young for wheat and barley market prices in Britain.

In recent years, abundant wheat harvests in the EC have resulted in huge commercial surpluses, thus exerting downward pressure on market prices which have often fallen below the intervention prices since 1984 (Figures 1 and 2). Consequently, not only will market prices be expected to fall, but also the transmission coefficient β will be expected to change value when such a situation occurs. As formulated, the price transmission model does not capture this phenomenon very well, since wheat market prices are supposed to fluctuate only between the bounds defined by the two policy prices. To reflect the fact that policies will not be fully effective under severe surpluses, the conceptual model is modified in two ways. Firstly, the intercept is hypothesized to be a linear, decreasing function of the variable, X_t; and secondly, the change in value of the transmission parameter β is taken into account by introducing a slope dummy variable into (3). As a result, expression (3) can be rewritten as:

(5) $$WP_t = [\theta - \theta_1 X_t] + [\beta_1 + \beta_2 DUM84]\,[PTH_t + k(X_t)\,(PIN_t - PTH_t)]$$

where DUM84 takes zero values before 1984 and 1 afterwards.[2] This more general formulation of the linkage between policy and market prices for wheat serves as a basis for the empirical analysis reported in this paper.

2. Preliminary graphical analysis has shown that market wheat prices started to fall below the intervention prices from the crop year 1984-85 onwards. For this reason, this year was selected as the data point at which the transmission coefficients are assumed to change values.

Model Implementation:

In order to estimate expression (5), it is necessary first to adopt an appropriate mathematical function for k(.), and to define the variable X_t. Both choices are necessarily limited by theoretical requirements and data availability.

One functional form that conforms to the maintained properties of k(.) is the logistic function given by:

$$(6) \qquad k(.) = \frac{1}{1 + EXP(\alpha_1 - \alpha_2 X_t)}$$

where: EXP designates the exponential function; and k(.) tends asymptotically to one when X_t is very large (abundant supplies) and towards zero in the opposite case. Between these two values, k(.) has a regular S-shape. Imposition of symmetry, of course, rules out other forms of behaviour.[3]

As far as the definition of X_t is concerned, its selection depends upon the time frame adopted - annual or monthly - to estimate the general price linkage equation. In the former case, a simple proxy is employed that is the percentage deviation of the total production of wheat relative to an historical exponential trend line:[4]

$$(7) \qquad X_t = QWD_t = \log(QW_t) - \tau_1 - \tau_2 t$$

where: QW_t is the total production of wheat in period t; t is a time trend

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3. An asymmetric response can be tested by incorporating quadratic or cubic terms of the variable, X_t, in the logistic function. In addition, it should be pointed out that in order to avoid estimation difficulties, no disturbance term has been added to equation (5). If it were stochastic, this would have led to the specification of a final price linkage equation characterized by a non-linearity in the price transmission coefficient and heteroscedastic error terms. This latter feature of the econometric equation requires a generalized least squares estimation approach (Raj and Ullah).

4. Coefficients τ_1 and τ_2 have been estimated from 1968 to 1985 for France and the Federal Republic of Germany, and from 1973 to 1985 for the United Kingdom.

which takes the following values: 1 in 196, 2 in 1969, and so on; and log designates a logarithm function.

The use of QWD_t as a proxy for X_t is justified on the grounds that under general economic, climatic and agronomic conditions, an expected "normal" harvest is captured by the exponential trend in the above equation and predetermines what is needed to balance the local supply with the corresponding demand components. Any deviation of wheat production from this trend generates a disequilibrium situation which, in turn, influences the local prices of wheat. Thus, a negative value of QWD_t results in an increased deficit, supplied by a larger volume of imports, and induces the market price of wheat to increase towards the threshold price. The positive value of QWD_t has the opposite effect and leads to market prices approaching the intervention price.

With a monthly time-frame, the local supply/demand conditions are captured by two variables which are linked to X_t through the following relationship:

(8) $$X_t = \delta_1 STW_t + \delta_2 NT_t \qquad \text{with } \delta_1 \text{ and } \delta_2 > 0$$

where: STW_t is the level of local stocks at the beginning of each month; and NT_t is the measured difference between exports and imports in each month.

The influence of these two variables on the market price of wheat is similar to that of the variable, QDW_t, the only difference being that the intra-year pattern of market prices is taken into account. This pattern is assumed to follow a seasonal cycle characterized by low levels at the beginning of each crop year when domestic supplies are plentiful and higher levels at the end of the crop year. In between, market prices should increase gradually according to the depletion of domestic stocks and the levels of net trade. Of course, it is understood that the magnitude of this yearly cycle will vary over time and differ for each EC country, depending in part upon the size of of the local annual wheat harvest.

Substitution of (6) for the function k(.) in (5) and adoption of the annual/monthly proxies (7) and (8) for X_t yields the following two empirical models:

Annual Time Frame

$$(9)\quad WP_t = \theta - \theta_1 QWD_t - \theta_2 DUM76 + [\beta_1 + \beta_2 DUM84A]\ [PTH_t + \frac{PIN_t - PTH_t}{1 + EXP(\alpha_1 - \alpha_2 QWD_t)}]$$

Monthly Time Frame:

$$(10)\quad WP_t = \theta - \mu_1 SWT_t - \mu_2 NT_t + [\beta_1 + \beta_2 DUM84M]\ [PTH_t + \frac{PINT_t - PTH_t}{1 + EXP(\alpha_1 - \sigma_1 SWT_t - \sigma_2 NT_t)}]$$

where $\mu_1 = \theta_1\delta_1$, $\mu_2 = \theta_2\delta_2$, $\sigma_1 = \alpha_2\delta_1$, $\sigma_2 = \alpha_2\delta_2$.

Note that both empirical models include dummy variables which account for the influence of institutional factors and structural change on the wheat market. The dummy variable, DUM76, allows for the impact of the SILO system. This policy regulation is expected to have a positive effect. Expressions (9) and (10) contain a slope dummy which captures the effect of unusually large wheat harvests during the crop years 1984 to 1985.

Econometric Results:

Expression (9) is estimated, using non-linear least squares (NLS) estimation techniques, for France and the Federal Republic of Germany over a sample period from 1968 to 1985, and for the United Kingdom for the period from 1973 to 1985. Due to limited data series, the monthly logistic model is only estimated by NLS for France over the period 1978.9 to 1986.7, and for the United Kingdom over the period 1981.9 to 1985.4.[5]

However, to take into account the influence of large wheat harvests on the values of the estimated coefficients, three different model specifications are estimated. The first one (models [A] or [AM]) deals with the case where the

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5. All data for the wheat prices and other regressors have been obtained from national sources and annual reports of the Statistical Office of the EC Commission. In the case of annual logistic models, all prices - market and policy - are averages computed on a crop-year basis.

logistic price transmission model is estimated using annual and monthly data samples, ending for the crop year 1983-84 (see Tables 1 and 2). The second econometric specification (models [B] or [BM]) is for the whole data sample, including the time period from 1984 to 1986 during which large wheat harvests were recorded. Finally, the last econometric specification (models [C] or [CM]) is similar to cases [B] or [BM], the only difference being that a slope dummy for the transmission coefficient has been incorporated into the logistic model.[6]

Dealing first with the annual models (Table 1), an inspection of the econometric results reveals that local supply/demand conditions (proxied by QWD_t) have partly influenced market wheat prices over the period and regions analyzed. Due to poor econometric performance, the original econometric specification was simplified as one in which the coefficient θ_1 was constrained to zero. With this alternative specification, it seems clear that the logistic model was suitable for the United Kingdom and the Federal Republic of Germany. In the case of France, the null hypothesis of zero value for the coefficient α_2 cannot be rejected for all three models. Based on these results, we can conclude that a constant price transmission model might be better-suited to explain the formation of average market wheat prices in France.

Regarding the monthly logistic model, the complete specification, as given by expression (10), was rejected on the basis of preliminary results and was replaced by a "hybrid" one in which some regressors were constrained to zero. Thus, the net trade variable NT_t is not incorporated in the French equation and only enters the British specification through the logistic response function. The variable for available supplies STW_t influences the French market price of

6. The logistic models for France have been estimated assuming that the structural change took place one year earlier during the crop year 1983-84. The reason for this is that the estimation of the initial annual model specification produces no econometric results due to the non-convergence of the non-linear least squares objective function.

wheat through the logistic term and the British one through the deterministic varying-parameter intercept.

An inspection of the econometric results contained in Table 2 suggests that as for the annual timeframe, a constant price transmission model would be preferable for France. In fact, the estimated coefficient σ_2 in the logistic function k(.) is not significantly different from zero for a level of significance of 5% in models [AM] nor [BM] and has the wrong sign in [CM]. In contrast, the estimated price transmission equation for Britain yields satisfactory results in terms of estimated coefficients and its explanatory power. However, these conclusions must be qualified due to the fact that several monthly specifications for France (models [BM] and [CM]) and Britain (model [BM]) are characterized by serial correlation among the estimated residuals (low DW statistic). In such circumstances, the estimated coefficients are inefficient and cannot be used for proper hypothesis testing (Johnston).

Varied findings concerning the imperfect transmission of policy prices can be inferred from the statistically valid estimated models. With the exception of one econometric specification (model [CM] for the Britain), the estimated parameters associated with the slope dummies are statistically significant and negative, thus implying not only a lower value of the price transmission coefficient, but also the presence of a structural change in the transmission of policy prices after 1984.

An application of a one-sided "t" test leads to mixed conclusions which differ among countries and model time frames. Concerning the annual models, the statistical results suggest that imperfect transmission of policy prices is prevalent in the German and British wheat sectors. Hence, in the former case, the null hypothesis that intervention and threshold prices are perfectly transmissible (β=1) is rejected at a 10% significance level for all three models. For Britian, only model [A] does not reject the null hypothesis of perfect transmission of policy prices. However, this latter result should be intepreted with caution in the light of multicollinearity problems associated with this econometric specification (characterized by a high condition number equal to 296), which yields an unstable variance/covariance matrix of the estimated parameters.

Conflicting results on the transmission of wheat policy prices can be derived from the three annual models estimated for France. In fact, comparing the French models [B] and [C] in Table 1, we observe that the former has a price transmission coefficient that is significantly different from one, whereas the opposite situation prevails for the later specification. This discrepancy can be explained by the inclusion of a slope dummy variable which, naturally, influences the values of the estimated coefficients on the one hand and, on the other, the inference drawn from statistical tests. Based on these considerations, and using only models [A] and [C], we conclude that policy prices for wheat in France are perfectly transmissible.

The "t" tests performed for the appropriate monthly logistic models in France (model [AM]) and Britain (models [AM] and [CM]) result in conclusions opposite in most cases to those found for the annual model. This lack of consistency is not surprising due to the use of different model specifications and the time aggregation factor in the annual case.

Finally, based on the above estimated models, it is interesting to examine the kind of "blend" policy prices to which EC regional market prices respond. For this purpose, a graphical representation of some estimated annual and monthly logistic response functions are displayed in Figures 3 and 4. Looking at the annual cases, it can be observed that the response functions k(.) are somewhat steep, suggesting that wheat market prices are very sensitive to changes in market supply/demand conditions. Thus, if a normal harvest has occurred for wheat in the three countries ($[QWD_t = 0]$), the value of k(.) gravitates around 0.95. This implies that under "normal" conditions, the market price of wheat in these three countries responds to a policy-determined price made up of 95% of the intervention price and the remaining 0.05% depending upon the threshold price. However, a shortfall of 20% to 30% in wheat production in the three countries induces a cereal market price response that depends totally upon the threshold prices.

An inspection of Figure 4 indicates a much more gradual response by the monthly logistic function to local supply/demand conditions than in the case of annual models. Thus, it is found that, when the British wheat market is in a situation of self-sufficiency (net trade equal to zero), the market prices respond to a blend policy price comprised of 30% threshold and 70% intervention prices. They will depend exclusively upon the intervention (k=1) and threshold (k=0) prices when monthly net trade is equal to -100,000 and 200,000 tonnes, respectively.

Conclusions and Policy Implications:

Despite varied and mixed estimation results, the application of a logistic-varying parameter model to represent the transmission of policy prices in the EC cereal sector seems promising. For two countries (the Federal Republic of Germany and Britain) out of three, we have not been able to reject this general model specification. Also there is evidence that unusually large wheat harvests have taken place in recent years, causing a "structural change" in the transmission of the policy prices. However, the use of different data samples and time-frames produced conflicting statistical results and inferences. For this reason, it would be useful to test the logistic-varying parameter model, using longer monthly time series, and to test this specification on other EC countries and cereals.

It is in the field of policy evaluation that the use of this specification could be fruitful. More specifically, this model could be helpful in quantifying the price effects on regional EC wheat markets of alternative policy measures aimed at reducing EC cereal surpluses (for instance, co-responsibility levies or set-aside program). Under such alternative policy regimes, due to an expected fall in wheat supply, a likely scenario is that the surplus position of some EC countries might be reduced significantly, which would imply that local market wheat prices would be driven up towards the threshold price. Such effects could be captured in the logistic varying-parameter model through the response of market wheat prices to the variable, QWD_t, and the stock and net trade variables in the annual and monthly specifications.

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TABLE 1: ESTIMATES OF THE ANNUAL POLICY PRICE TRANSMISSION MODEL FOR FRANCE, THE FEDERAL REPUBLIC OF GERMANY AND BRITAIN

Country	Model	Intercept	Dummy Variable (DUM76)	Transmission Coefficient	Slope Dummy DUM83A or DUM84A	Weight a_t Parameters of the Logistic Function k(.)		R^2	DW	COND
		θ	θ_2	β_1	β_2	α_1	α_2			
France	[A] (1968-1982)	14.1554 (0.25)	114.676 (2.92)	0.956327 (10.81)		-8.37786 (-1.20)	49.4765 (1.17)	0.988	1.45	34.5
	[B] (1968-1985)	119.313 (3.28)	169.5 (5.08)	0.78183** (13.5)		-54.9809 (-0.00)	330.311 (-0.00)	0.988	1.34	16.4
	[C] (1968-1985)	8.949 (0.16)	111.704 (2.81)	0.96468 (10.75)	-0.06962 (-2.51)	-8.48487 (-1.21)	50.4034 (1.18)	0.991	1.44	36.6
Federal Rep. of Germany	[A] (1968-1983)	71.0608 (2.02)	48.5721 (5.67)	0.85258* (9.74)		-3.17042 (-2.16)	21.1799 (2.05)	0.979	1.71	51.6
	[B] (1968-1985)	197.374 (2.02)	25.0233 (1.23)	0.530715** (2.24)		-0.412444 (-0.57)	42.3880 (0.99)	0.932	1.37	81.0
	[C] (1968-1985)	46.1456 (1.99)	50.0885 (6.86)	0.865443* (6.86)	-0.12412** (-7.35)	-3.66107 (-2.73)	24.7759 (2.42)	0.983	2.16	51.2
United Kingdom	[A] (1973-1983)	27.9111 (1.08)	13.5002 (4.27)	0.556441 (0.99)		1.77579 (0.08)	3.54858 (0.10)	0.997	1.91	296
	[B] (1973-1985)	12.3771 (3.83)	10.4039 (3.73)	0.868268** (16.32)		-2.16502 (-3.33)	13.4778 (3.11)	0.990	2.25	21.0
	[C] (1973-1985)	30.1794 (21.02)	13.2648 (8.38)	0.521402** (28.30)	0.05749** (5.37)	10.1313 (0.69)	44.0397 (0.69)	0.998	1.86	175.0

DW = Durbin-Watson statistic, R^2 = coefficient of determination, COND = Condition number and the numbers in parentheses are asymptotic "t" values

(*) and (**) indicate that the transmission coefficients β are statistically different from one for a 10% and 5% levels of significance, respectively.

TABLE 2: ESTIMATES OF THE MONTHLY POLICY PRICE TRANSMISSION MODEL FOR FRANCE AND BRITAIN

Country	Model	Intercept	Beginning Stocks (SW_t)	Transmission Coefficient		Weight a_t — Parameters of the Logistic Function k(.)		R^2	DW	COND
					Slope Dummy DUM83A or DUM84A					
		θ	μ_1	β_1	β_2	a_1	σ_2			
France	[AM] (1978.9-1983.7)	137.328 (7.59)		0.83241** (27.33)		0.64214 (2.09)	0.01259 (1.21)	0.987	1.60	38.6
	[BM] (1978.9-1986.7)	381.132 (9.69)		0.65279 (13.5)		0.06581 (0.09)	330.311 (1.48)	0.988	0.33	28.2
	[CM] (1978.9-1986.7)	228.482 (4.88)		0.69051 (19.25)	-0.09553 (-7.18)	0.68725 (0.90)	-0.33727 (-2.92)	0.894	0.54	23.9
United Kingdom	[AM] (1981.9-1984.7)	-6.40254 (-0.50)	-0.85336 (-3.94)	0.88616 (10.31)		1.1473 (1.51)	3.30531 (6.29)	0.875	1.12	51.5
	[BM] (1981.9-1985.4)	29.0265 (2.03)	-1.23402 (-6.59)	0.73976** (6.23)		-1.17157 (-6.59)	13.4778 (5.40)	0.823	0.84	54.9
	[CM] (1981.9-1985.4)	-4.16378 (-0.27)	-0.73429 (-2.10)	0.89851 (9.88)	-0.0876 (-5.14)	0.58848 (0.37)	3.45991 (3.22)	0.891	1.33	66.6

DW = Durbin-Watson statistic, R^2= Coefficient of determination, COND= Condition number and numbers in parentheses are asymptotic "t" values.

(**) indicates that the transmission coefficients β are statistically different from one for a 5% level of significance,respectively. Note that no statistical test has been performed on the monthly logistic specifications (models in France (models [BM] and [CM]) and Britain (model [BM]). The reason for this is that, in all these models, coefficient estimates are inefficient due to the presence of serial correlation among the estimated residuals.

FIGURE 1: PRICE OF WHEAT IN FRANCE INTERVENTION, THRESHOLD AND MARKET PRICES,

AUGUST TO JULY CROP YEAR

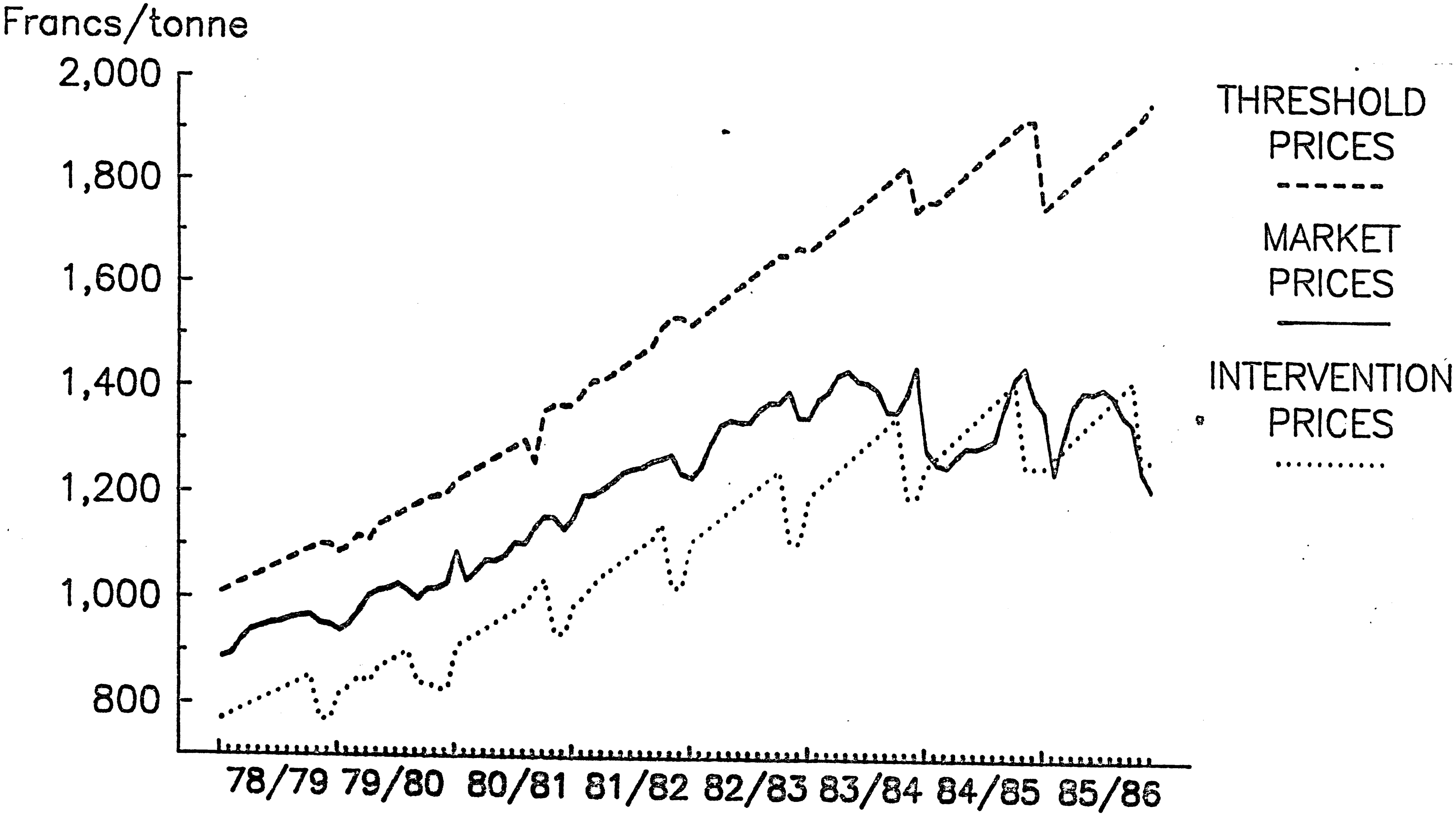

FIGURE 2: PRICE OF WHEAT IN BRITAIN INTERVENTION, THRESHOLD AND MARKET PRICES, AUGUST TO JULY CROP YEAR

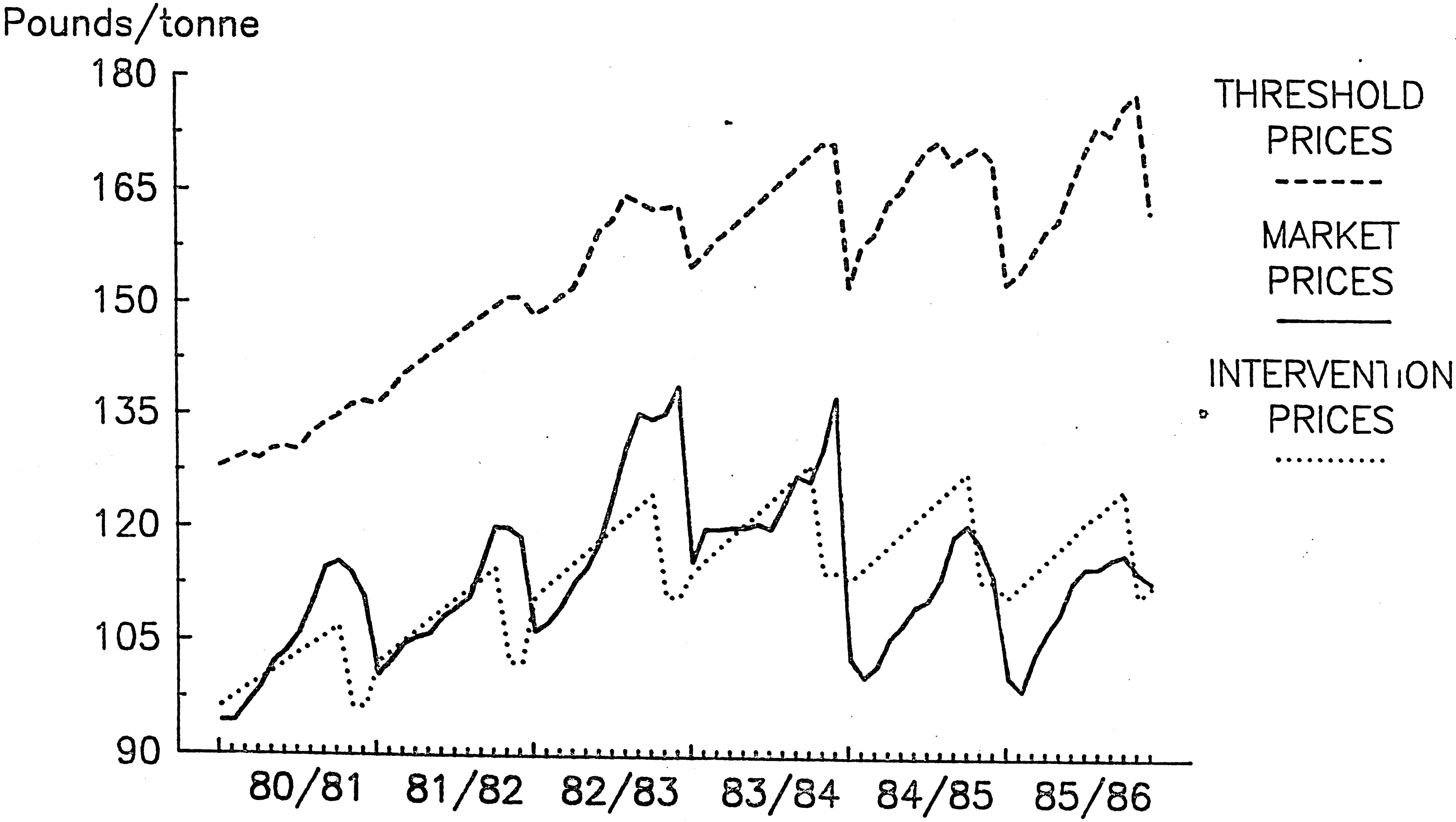

FIGURE 3: LOGISTIC RESPONSE FUNCTION k(.)

BASED ON THE ESTIMATED ANNUAL MODEL SPECIFICATIONS

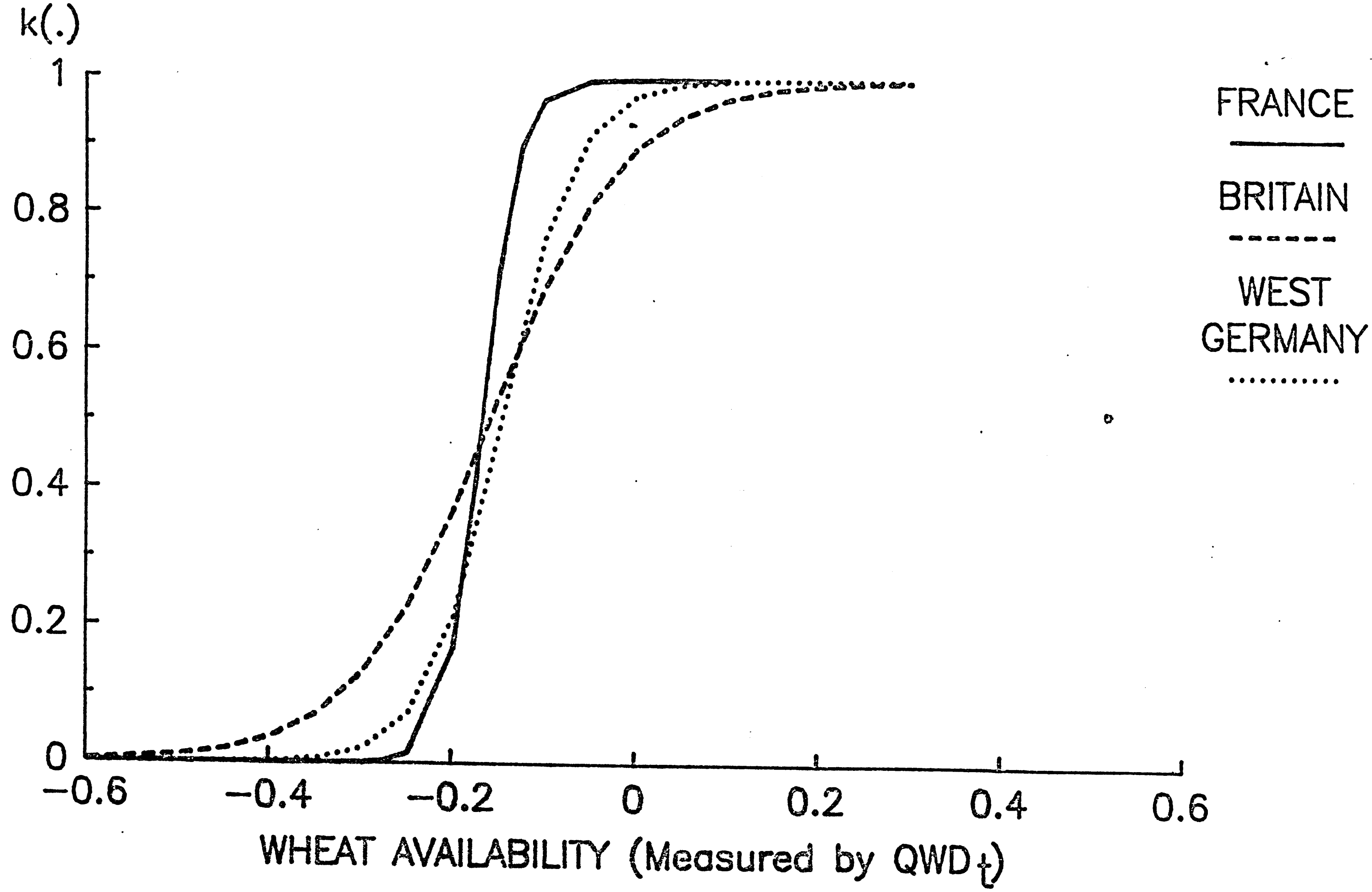

FIGURE 4: LOGISTIC RESPONSE FUNCTION k(.) FOR BRITAIN

BASED ON THE MONTHLY MODEL (C)