A Cointegration Analysis of the Long-Run Supply Response of Spanish Agriculture to the Common Agricultural Policy*

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Abstract

Using cointegration techniques, we estimate two models that capture the long-term relationship between Spanish prices and agricultural production. The models were estimated over Spanish agricultural data from 1970 to 2000, a period spanning Spain's implementation of the Common Agricultural Policy in 1986 and the application of the MacSharry Reforms in 1992. The models, as well as plausible counterfactual scenarios constructed to assess the production changes induced by the CAP, lead to three principal results. First, we find that Spanish agricultural output is responsive to agricultural prices. Second, we find that the MacSharry reforms have been instrumental in restraining agricultural ~~production. Third, we find that Spanish agricultural output would have been higher if Spain had not~~ applied the CAP. These results are important and have broad implications. First, they strengthen the position of those reformers both within and outside of Europe that argue for lower price supports as an appropriate policy for stemming European agricultural surpluses. Second, they indicate that recent EU reforms, which have in effect extended the MacSharry reforms, are appropriate measures for curbing European agricultural surpluses.

Keywords: Price Support Policy; MacSharry Reforms; Cointegration Techniques; European Economic Integration.

JEL Codes: Q11, Q18, C32, F02.

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I. INTRODUCTION

The supply response of European agriculture is a source of controversy both among policy makers and academicians. Since the inception of the Common Agricultural Policy (CAP), European Union (EU) agricultural production has expanded dramatically, exceeding domestic requirements and placing financial preassure on the EU budget.[1]

European policy makers do not dispute that they face a problem with surpluses, but they disagree as to the cause. One group contends that domestic support programs have encouraged European farmers to overproduce, causing a disruption of world trade in agricultural products and an inefficient allocation of resources worldwide.[2] The other group retorts that price support programs have had little impact on generating these surpluses because European agriculture is insensitive to price changes; instead, they assert that the surpluses are caused by the supply-enhancing effects of technological improvements such as the use of new or better seeds, fertilizers and cropping methods. Perhaps not surprisingly, this disagreement as to the causes of these surpluses also extends to the policy arena. For instance, there still exists disagreement regarding the effectiveness of the

[1] For brevity, the term EU stands for the European Communities, the European Community and the European Union throughout the paper.

[2] In a speech before the European Parliament, on May 19, 2003, Franz Fischler, then EU Agriculture Commisioner, states: "I know that there are those of you that question the necessity and extent of our market reforms ... [But our] priorities are clear. We want better market access for all, a reduction in trade-distorting farm subsidies and all forms of export aid...".

MacSharry reforms initiated in 1992, a set of comprehensive reforms that incorporated constraints on production as well as reductions in price supports.

Surprisingly, few empirical studies have been marshaled by either side to buttress their respective arguments. The *lone exception* is a study by Bouchet, Orden, and Norton (1989) who, using French agricultural data from 1959-84, find that supply was indeed price inelastic, changing little with price changes.

However, the study had two major limitations. First, during the entire time period covered by the study, France had in place an extensive farm program that raised prices and limited their variability. That is, there were no observations corresponding to a period when support programs were not in place.[3] Second, and more importantly, the estimation techniques that they employed did not allow for testing and estimating the long-term dynamic relationship between prices and production.[4]

The broad objective of this paper is to address the void in the empirical literature on the long-term relation between CAP policies -- the price support system and the MacSharry reforms -- and European agricultural production.

[3] One could of course argue that the CAP did not actually go into effect in 1968. However, policies before and after the CAP were very similar. The goal of French agricultural policy in the immediate post-war period was to eliminate food shortages at any cost and to do so largely with high, guaranteed price supports. By 1950, the government had achieved its objective of adequate food supplies and surpluses became evident in wheat, wine and sugar-beet production. The goverment was pressured by farmers to replace the already generous Monnet Plan for 1947-1950 with the Second Plan for Modernization and Equipment that provided for a further increase in agricultural production of 20 percent. At the time of the Third and Fourth Plan (1957-65), the government appeared resigned to the fact the additional production was 'inevitable' and sought to eliminate surpluses through exports. The entire period prior the CAP was characterized by agricultural surpluses in France.

[4] In fairness to the authors, appropiate techniques, such as cointegration analysis for integrated variables, were in their infancy at the time.

Specifically, we do so by applying cointegration techniques to the production response of Spanish agriculture to the CAP from 1970 to 2000. Spain provides an ideal case to explore the impact of these policies since its important agricultural sector witnessed a radical transformation during this thirty-year period both in government policy and in access to the European market. Prior to entry in the EU in 1986, agricultural policy in Spain provided a relatively low level of protection for farmers and Spanish farm products enjoyed only limited access to the EU. Following a transition period, support prices and market regulations became identical to those in the EU and free access to the European market was attained by 1993.[5]

The cointegration techniques that we use capture the long-term relation between prices and production by exploiting the time-series properties of the data set. This permits us to create three counterfactual scenarios to assess the agricultural production shifts brought about by the implementation of the CAP price supports and the MacSharry reforms of 1992. The three cases share the assumption that Spain does not implement the CAP, including the MacSharry reforms, but they assume distinct institutional arrangements under which expected world agricultural prices evolve. In the first scenario, government agricultural

[5] Although rigid controls existed in some markets, most Spanish products either lacked support or support levels were little above world prices before entry in the early 1980s. An early calculation, EC (1987, p. 17), placed the weighted average price increase from replacing Spanish support prices with CAP prices at 14.1 percent. For further evidence on support price differences, see Rollo (1980), FAO (1980), USDA (1983), and Tio (1986). In addition, exports to the EU were only 19.3% of total production at the time. Between 1984 and 1992, intervention prices were increased to CAP intervention levels and market regulations became identical to those in the Common Market Organizations. By 2001, exports to European markets had reached 38.5% of total production.

policies worldwide remain unchanged so that Spanish internal prices are above world prices by a nominal tariff rate consistent with previous Spanish trade policy. In the second scenario, Spain is assumed to move unilaterally to free trade in agricultural products. Finally, in the third scenario, all world trade barriers to agricultural products are removed and Spanish internal prices equal world prices in a liberalized world agricultural market.

The paper contains four Sections. Section II is devoted to the description of the agricultural policy in Spain since the early seventies. The econometric model is described in Section III, and results are presented in Section IV. Section V offers concluding comments.

II. THE EVOLUTION OF AGRICULTURAL POLICY IN SPAIN

In this Section we give a brief description of the evolution of the agricultural policy in Spain since the early seventies, focussing on the implications of EU membership and the 1992 reforms of the CAP for Spanish agriculture.[6]

II.1 Agricultural Policy Change in Spain

The CAP and its reforms work principally by using changes in prices to induce changes in agricultural supply.[7] Early in the eighties, Spanish support

[6] Salmon (1991), Barceló and García-Alvarez-Coque (1987), and Hine (1989) contain excellent analyses of the implications of EU membership for Spain. See also Rollo (1980), USDA (1983), and Tió (1986) for other good descriptive studies of Spanish agricultural policies at the time of accession to the EU. A detailed description of the 1992 reforms can be found in Swinbank (1993 and 1997).

[7] Indeed, it was the prospects of higher agricultural prices (and the consequences for farm output, employment and income) that played one of the key roles in Spain's decision to apply for EC membership in July 1977. At the time, agriculture played a dominant role in the Spanish economy,

prices were close to world prices as Spain had progressively reduced effective rates of protection during the 1970s (see MAPA, several issues). Thus, the CAP's application meant that Spanish farmers would face support prices that were far more generous and that applied to a much broader range of goods.[8]

Concern over the budgetary effects of extending the CAP to Spain, given its large agricultural sector and lower guarantee prices as well as the political pressure from those EU farmers likely to face heightened competition, led existing EU members to impose a transition period on Spanish products. For most products, the transition period was set at seven years, during which time Spanish support prices would be raised annually by 1/7th of the price difference that existed between Spanish and EU support prices during the 1985/86 crop cycle.[9] Prices would converge immediately for products whose price differentials were smaller than 3 percent. For the few products in which Spanish support prices were higher, such as sugar beet and dairy products, Spanish prices would be frozen and convergence eventually achieved via the increase in EU prices.

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accounting for 23 percent of total employment and almost 10 percent of GDP. Agriculture was also the principal source of employment and income for Spain´s poor.

[8] In addition to price support systems, prior to entry Spanish farmers received subsidies aimed largely at increasing agricultural productivity. Most of these subsidies continued, under renamed programs, after entry into the EU, but their relative importance declined significantly.

[9] Fruits and vegetables were initially given a ten-year transition period consisting of two stages. For the first four years, Spanish prices were frozen in order to allow farmers elsewhere in the EU to prepare for increased competition from Spanish products. This was to be followed by a six-year period during which prices would be raised by 1/6th of the existing price differential so as to attain convergence by the end of 1995. Olive and other vegetable oils also received a similar ten-year transition period, except that during the first five years olive oil prices would be increased annually by 5 percent of the difference between Spanish and EU prices, whereas vegetable oil prices would experience a 'standstill', i.e., the status quo would remain unchanged. Then for the remaining five years, prices for both would rise by equal annual increments. Eventually, as integration speeded up, transition periods for most products actually ended by January, 1993.

II.2 1992 CAP Reforms

EU budgetary pressures in the operation of the CAP during the eighties led to a series of reforms culminating in the so-called MacSharry reforms of 1992. These reforms sought to relieve pressure on the EU budget by lowering support prices and, to a lesser degree, by introducing supply control measures. A new, but pivotal, element of the reforms was the introduction of monetary transfers designed to decouple income support to farmers from production levels.

The changes were substantive in three sectors: cereals, oilseeds, and beef and sheepmeat. The cereal support price was reduced by 29 percent to be phased in over three years beginning with the 1993 crop year. Guaranteed prices for oilseeds and protein crops were eliminated. To compensate for the revenue loss from these price changes, farmers would receive direct payments based in part on historical acreage and yields, with large farmers being required to set aside 15 percent of their acreage to be eligible for the payment.[10] Beef prices were lowered 15 percent over three years while sheepmeat prices were frozen at existing levels. To offset the income loss, a headage payment limited by historical stocks and quotas at national level was introduced.[11]

III. THE MODEL

Underlying our empirical analysis is a simple supply and demand model of the Spanish agricultural sector. Below, we first outline the basic theoretical

[10] In the case of cereals, this historical acreage was defined as the average of the cultivated land in the years 1989 to 1991.
[11] For beef products, stocking rate limitations were also required.

framework and the impact of the policy changes reviewed in the previous section under this framework. Next, we propose an econometric model which permits us to identify the dynamic reduced-form empirical relations that are explored in the next section.

III. 1 A basic model of agricultural supply and demand

We begin by assuming that the Spanish agricultural sector is divided into corporate farms and family farms. The former operate as profit-maximizing price takers who produce output Q_c, using intermediate inputs M_c purchased at price P_M, and the services of land, V_c, capital, K_c, and labor, L_c, obtained at market rates R_V, R_K, and W, respectively. Agricultural output is sold at given price P. Family farms produce output Q_f and also employ intermediate inputs M_f, land, V_f, capital, K_f, and labor, L_f. Although inputs are also bought at price P_M and land and capital services are purchased at market rates R_V and R_K, the hourly wage of family workers is determined as a residual. Total agricultural output is the sum of the output in the two sectors, $Q_c + Q_f$.

All agricultural producers behave as price-takers, i.e. as if facing a perfectly elastic demand curve for the product. For those producers under a guaranteed price umbrella, the government purchases excess supply and the market clears. For producers without a guaranteed price, world prices prevail in the domestic market, which is cleared by the flow of foreign trade. Agricultural producers are also price takers in input markets.

III.2 Impact of EU entry and the 1992 CAP reform

For most Spanish producers, entry into the EU represents both an increase in agricultural prices as well as a reduction in input prices. Product price increases resulted from two upward demand shifts. First, a policy-induced demand shift due to increases in existing guaranteed prices for some products and the implementation of new support price regimes for other products. Second, a market-induced demand shift resulting from an increase in foreign demand as Spanish exports, such as fruits and vegetables, gained preferred access to EU markets. The ensuing higher prices in turn brought about increases in quantities supplied as producers moved upward along their supply curves. Output also rose as the reduction in real input prices (stemming from the increased competition in the Spanish market from EU suppliers) shifted agricultural supply outwards.[12]

The impact of the 1992 reforms is to some extent less clear. The reduction in the level of support prices would generally lead to a fall in input usage and farm output. However, two additional elements of the scheme, one designed to further reduce excess supplies, and the other designed to compensate for revenue losses stemming from the reduction in intervention prices without stimulating production, may not work as intended. First, the set-aside could be non-binding

[12] Price support programs cause an additional outward shift in supply since they also decrease the variance of prices throughout the season which in turn reduces the riskiness of farm operations. While not denying the importance of this effect, we do not attempt to isolate its magnitude from the overall price effect. Since some subsidies were designed to facilitate the acquisition of new technologies, we will assume that their effect is also embedded in the effect of overall technological improvements. For a recent survey of theoretical analyses of agricultural policies, see Bullock and Salhofer (2003). For an early discussion of the effects on Spanish farmer's income of the reduction in real input prices following entry, see San Juan (1995).

and, thus, have no impact on output, especially in the long run. Note that as the optimal land use falls with decreasing intervention prices, this decline in cultivated land may exceed the amount required by the set-aside restriction. Additionally, since set-aside land could still be used for some crops, overall agricultural output could actually increase.[13]

Second, the 1992 reforms attempted to compensate farmers with direct payments based on historical land use. These payments, while purportedly designed to be decoupled from production, are effectively a subsidy to land that encourages a more intensive use of the historical acreage. This is so because, in equilibrium, the land market price net of compensatory payments decreases for initial land owners, and farmers experience an income effect which will lead them to use other inputs more intensively resulting in higher output.[14]

III. 3 The Statistical Model

As in Granger (1984) and Engle and Granger (1987), consider the following general definitions useful in developing the statistical model. Let $\mathbf{z}_t$ be the m x 1 vector at time t of $I(d)$ variables.[15] If a linear combination, $\gamma'\mathbf{z}_t$, is of order $I(d - b)$, b >0, then the series are said to be cointegrated $CI(d, b)$ and the vector of coefficients γ

[13] Compulsory set-aside land may be used to produce materials for the manufacture of products not intended for human or animal consumption. In the case of voluntary set asides, this restriction only applies to products intended for human consumption.
[14] Although it may seem a paradox that the use of inputs other than land increase while land use either increases or decreases, note that land use is limited by historical acreage, so no substitution effect in the standard meaning is possible.
[15] Variables are integrated of order d, denoted by $I(d)$, if they are stationary in mean and variance only after differencing d times. Early general references to non-stationary time series include Box and Jenkins (1970) and Granger and Newbold (1986).

is called the cointegrating vector. For m variables, the maximum number of linearly independent cointegrating vectors is m - 1. If the series are cointegrated $CI(1,1)$, then there exists a vector autoregressive representation, VAR:

$$\boldsymbol{\Gamma}(L)\nabla \mathbf{z}_t = \mathbf{C}_0 + \mathbf{C}_1 t + \boldsymbol{\Pi}\, \mathbf{z}_{t-1} + \mathbf{e}_t, \qquad t = 1,2,... \tag{1}$$

where L is the lag operator, $\nabla = (1 - L)$, $\boldsymbol{\Gamma}(L)$ is a finite-order matrix lag polynomial, $\mathbf{C}_0$ and $\mathbf{C}_1$ are m x m matrixes, $\boldsymbol{\Pi}$ is an m x m matrix of rank $r < m$, which permits its decomposition into two full rank matrices $\boldsymbol{\Pi} = \mathbf{A}_{m\text{x}r}\mathbf{B}_{r\text{x}m}$, and $\mathbf{e}_t$ has the usual white noise properties. Granger and Engle (1987) show that $\mathbf{B}_{r\text{x}m}\mathbf{z}_{t-1}$ are r stationary processes so that the rows of matrix $\mathbf{B}_{r\text{x}m}$ consist of r linearly independent cointegrating vectors, also called the cointegrating equations, CE. Moreover, since $\nabla \mathbf{z}_t$ depends (through matrix $\mathbf{A}_{m\text{x}r}$) on these stationary deviations from the cointegrating relations, Equation (1) can be seen as an error correction mechanism, ECM.

A key feature of Spanish agriculture is that the joint evolution of input and output prices, technology, and family labor is such that we cannot reject the hypothesis that all variables are weakly exogenous with respect to output with the exception of land services.[16] This permits us to employ a restricted version of Equation (1). First, define $\mathbf{x}_t = (p_t, p_{Mt}, w_t, r_{Kt}, l_{ft}, A_t)'$ as the 6x1 column vector at time t of prices, family labor, total factor productivity, and subsidies where lower case letters represent the logarithmic transformation of the original variables. Next,

[16] See the next Section for the reported results of the test and Engle (1984) for its description.

denote $\mathbf{y}_t = (q_t, r_{Vt})'$ and $\mathbf{z}_t = (q_t, \mathbf{x}_t')'$. If column vector $\mathbf{z}_t$ is cointegrated of order $CI(1,1)$ and the elements of $\mathbf{x}_t$ are exogenous $I(1)$, then there is a sub-system VAR for $\nabla \mathbf{y}_t$:[17]

$$\nabla \mathbf{y}_t = \mathbf{c}_0 + \mathbf{c}_1 t + \lambda \nabla \mathbf{x}_t + \Sigma_i \Psi_i \nabla \mathbf{z}_{t-i} + \mathbf{\Pi}_1 \mathbf{z}_{t-1} + e_{1t} \quad (2)$$

where $\mathbf{c}_0$ and $\mathbf{c}_1$ are $2x1$ column vectors, λ and Ψ_i are $2x6$ matrices and $\mathbf{\Pi}_1$ is the corresponding submatrix of $\mathbf{\Pi}$ in Equation (2). This is a very flexible and powerful specification. The first two terms capture the effects of deterministic trends in output and land services. The third reflects the contemporaneous effect of changes in prices, technology and family labor, whereas the fourth has the lagged effects of all variables. $\mathbf{\Pi}_1$ can be motivated as long-term relations derived from a supply and demand system. Since exogenous variables drive the evolution of agricultural production in the long term, the production impact of price changes stemming from the introduction and reform of the CAP is permanent and does not die out.

Note that we include two stocks, total factor productivity and family labor in $\mathbf{x}_t$. Generally, total factor productivity would increase production and reduce prices over the long term. This is certainly the case for the global agricultural market. In Spain's case, this relation may exist before entry to the EU for those products without a price support program. However, after entry to the EU, technological improvements should continue to have a positive effect on

[17] In addition to the absence of correlation between shocks, weak exogeneity implies that price changes and technological improvements influence output changes contemporaneously and will be long-run forcing variables for output. See Pesaran et al. (2000) and Harbo et al. (1998) for details.

agricultural output, even though output prices are largely determined by policy makers.

We also include family labor to capture structural change within the agricultural sector in Spain. Surveys of the Spanish farm sector consistently show that the vast majority of small farms only employ family labor (see Censo Agrario, several issues). There is also evidence that the incomes received by these family workers entail a lower hourly wage than that received by hired workers, suggesting immobility and disguised unemployment (see Mora and San-Juan, 2002). In addition, the number of small farms as well as their share of total production has fallen steadily in the last thirty years in line with the decline of family labor in agriculture. Thus, over our sample period the family-farm sector exhibits a secular, unbroken decline independent of the evolution of prices of products and inputs.

Before leaving the section, it should be stressed that Equations (1) - (3) can be extended to include the effects of stationary variables. In our case, these include annual deviations from historical averages of rainfall and temperature, $Rain$ and $Temp$, compensatory payments, CP , and the rate of set aside, SA. As this does not imply a substantial change in the model, we have not included the stationary variables to avoid excessive notation.

IV. MODEL ESTIMATION AND RESULTS

Drawing on the time-series and cointegration literature, we proceed in a sequential fashion. First, we carry out unit root and cointegration tests to evaluate the statistical properties of the variables in our dataset and test for the existence of cointegrating relationships. As we are unable to reject that the series are integrated of order 1 and that they are jointly cointegrated, we then estimate an ECM model that captures both the short-term dynamics of the variable set and the long-term cointegrating relations. Lastly, we simulate the evolution of the system of variables under three counterfactual scenarios designed to assess the impact of the CAP and its reform on Spain's agricultural sector.

IV.1 Unit Root Tests

We employ annual data for the period 1970 to 2000. Variable definitions and data sources are listed in Table 1.[18] Graphical inspection of the evolution of the variables in levels and differences, as well as their estimated autocorrelation and partial autocorrelation functions suggests that the variables are, at most, integrated of order 2.

[18] See the Appendix for a detailed description of the data sources and the definitions of the variables.

Table 1: Variable Definitions and Sources.[a]

Variables	Source	Definition
Quantities:		
q	*Eurostat*	Index of Final Agricultural Production
l_f	*MAPYA*	Annual Work Units of Family Labor
Prices:		
p	*Eurostat*	Agricultural Production Fischer Price Index
p_M	*Eurostat*	Intermediate Input Fischer Price Index
r_V	*MAPYA*	Land Rental Rates
r_K	Ball *et al (2003)*	Capital Rental Rates
w	*INE*	Hourly Industry Wage
Other Factors:		
TFP	*USDA*	Index of Farm Productivity, 5-year Moving Average
Rain	*INM*	Annual Average Rainfall
Temp	*INM*	Annual Average Temperature
CP	*MAPYA*	Compensatory Payments
SA	*Eurostat*	Spanish Set Aside Rate

Notes: All variables are in logarithms and normalized at 1970 values except *Rain* and *Temp* , which are standardized annual deviations from the historical average and CP and SA which are normalized at 1993 values. See the Appendix for a detailed description of the variables.

[a] Eurostat stands for the Statistical Office of the EU; MAPYA is Spanish acronym for Spanish Ministry of Agriculture; INE stands for the Spanish National Statistical Office; USDA for the US Department of Agriculture and INM for the Spanish National Meteorological Institute.

Table 2 shows the results of the iterated testing procedure suggested by Dickey and Pantula (1987) for the Augmented Dickey-Fuller, ADF, test (Dickey and Fuller, 1979), the Dickey-Fuller test when additive outliers are present , DFAO, (see Vogelsang, 1999) and also the Phillips-Perron, PP, test of unit roots (Phillips and Perron, 1988). The test results for the null hypothesis of two unit roots are listed in the top panel of Table 2.

Table 2: Unit Root Tests: Augmented Dickey-Fuller (ADF), Dickey-Fuller with Additive Outliers (DFAO), and Phillips-Perron (PP) Tests.

Variable	ADF	DFAO	PP
H0: Variable in levels has 2 unit roots. H1: Variable in levels has 1 unit root.			
Δq	-4.80***	-6.86***	-7.56***
Δl_f	-3.80**	-7.91***	-5.85***
Δp	-4.49***	-5.02***	-5.08***
Δp_M	-4.02***	-2.37	-3.63**
Δr_V	-3.39**	-3.39**	-2.12
Δr_K	-3.98***	-3.69**	-3.58**
Δw	-3.35*	-4.10***	-3.63**
ΔTFP	-2.60	-5.35***	-5.38***
$\Delta Rain$	-5.87***	-7.49***	-8.66***
$\Delta Temp$	-5.03***	-4.97***	-8.84***
H0: Variable in levels has 1 unit root. H1: Variable in levels has no unit roots.			
q	-1.99	-1.99	-3.39*
l_f	-1.36	3.42	-1.25
p	-1.60	-0.85	-0.48
p_M	-1.75	-0.02	-0.26
r_V	-0.97	-1.79	-1.74
r_K	-2.15	-2.43	-1.72
w	-1.34	-2.30	-0.23
TFP	-2.20	-1.51	-1.61
$Rain$	-3.10**	-3.54**	-3.59***
$Temp$	-3.71**	-4.96***	-4.96***

Notes: Significant coefficients are indicated by *, **, ***, for significance at the 10%, 5% and 1% level, respectively. The Durbin-Watson *d*-statistic and the Ljung and Box (1978) Q* statistic were computed on the fitted residuals to tests for absence of first-order and serial correlation respectively. All variables were detrended except *Rain*, r_V, r_K , and k, which were demeaned.

Although the sample size is small and the alternative hypothesis differs to some extent for each test, the results convey a consistent message.

In all cases but in the price of land services the null is rejected at the 99 per cent level in at least one of the tests. In addition, the null of two unit roots is rejected in two tests at the 99 per cent level in ten out of thirteen cases. In the case of the price of land services, rejection with the ADF and the DFAO tests are very near the 99 per cent level, as the *p*-value for both tests is 98.9 percent.

The bottom panel of Table 2 reports the results of the tests of one unit root against the alternative of absence of unit roots. Except for *Rain* and *Temp* we are unable to reject the null hypothesis at the 90 per cent level for the ADF and the DFAO tests. We do reject the null for output using the PP tests, but we ignore these results since the test fails to take into account the high autocorrelation present in the variable.

To sum up, all the variables in our dataset, with the exception of *Rain* and *Temp*, satisfy the statistical properties necessary for the existence of cointegrating relations. The results for *Rain* and *Temp* should not be surprising and confirm our intuition that annual deviations from historical averages of rainfall and temperature show stationarity in mean and variance. In the following, weather conditions will be included as the ratio of *Rain* to *Temp* to simplify the variable specification.[19]

[19] We tried other specifications, including a quadratic polynomial for the two variables. Both the main estimation results and the simulations were unchanged.

IV.2 Cointegration Results

To test for the existence and number of cointegrating relations, we present in Table 3 the results of applying the Johansen (1991) procedure for a variety of models.[20]

We begin our exploration of long-term relations in the data without imposing weak exogeneity on nonstationary $I(1)$ variables as in Equation (1) with $\mathbf{z} = (q, \mathbf{x}')'$.[21] Given that all variables are in nominal terms, we expect the presence of a linear trend in levels reflecting, amongst other things, the differing impact of inflation on each of them. Two different specifications of the deterministic components of the series can account for the fact that a linear trend is present in the CE. The first specification includes a trend and a constant in the VAR for each variable of vector **z**. A quadratic deterministic trend in levels for a small sample may account for observed permanent changes in the slope, as could be the case for output after entry in the EU or the 1992 reforms. The second specification includes the trend in the CE, which implies rank restrictions on the deterministic parameters. We performed a likelihood ratio test on these restrictions and strongly rejected the null. Consequently, we proceed with the model with trend in the VAR.

[20] The procedure tests for the number of cointegrating vectors by assessing the rank of the ML estimate of matrix $\mathbf{\Pi}$. The test proceeds sequentially. First, the null hypothesis of absence of a cointegrating vector (i.e. H_0: $r = 0$) is tested. If rejected, the null of one cointegrating relation (i.e. H_0: $r = 1$) is then tested, and so on until a null can no longer be rejected. We computed both the max(λ) and the trace statistic. Both statistics have the same null hypothesis (H_0: rank($\mathbf{\Pi}$) = r) but the alternative for the former is H_1: rank($\mathbf{\Pi}$) = r +1, while for the latter it is H_1: rank($\mathbf{\Pi}$) > r.

[21] The residuals pass both the Multivariate Ljung-Box statistic and the Omnibus normality statistic at the 90% level when the number of lags in the VAR in levels is one. The contemporaneous and one-lag values of stationary exogenous variables *CP*, *SA*, and weather conditions were also initially included in the VAR specifications.

Table 3: Cointegration Tests and VAR Model Diagnostics

Panel A: Johansen Tests for the Rank of Cointegration

	Model 1		Model 2		Model 3		Model 4	
H_0: rank(**Π**)=r	**max(λ)**	**trace**	**max(λ)**	**trace**	**max(λ)**	**trace**	**max(λ)**	**trace**
$r = 0$	58.63**	131.3**	59.14**	176.6**	22.46**	22.96*	22.26**	31.55**
$r = 1$	35.20	72.69	45.91	117.4	0.499	0.499	9.282	9.282
$r = 2$	14.45	37.49	29.84	71.52	--	--	--	--

Panel B: Cointegrating Equation

	Model 1		Model 2		Model 3		Model 4	
	Coef.	$B_{1j}=0$	**Coef.**	$B_{1j}=0$	**Coef.**	$B_{1j}=0$	**Coef.**	$B_{1j}=0$
Weight in ECM	0.04		-0.04		-0.27		-0.27	
q	1.43		0.62	7.58**	3.28	85.75**	2.49	51.61**
p	-0.88	25.41**	-0.17	2.70*	-0.26	1.51	-0.20	2.92*
p_M	0.73	259.1**	-0.03	0.18	0.09	0.46	--	--
r_V	0.15	3.13*	0.26	108.3**	0.08	0.37	0.06	0.27
r_K	0.02	0.87	0.11	251.1**	--	--	--	--
w	0.01	0.02	-0.06	0.69	--	--	--	--
TFP	-2.87	29.93**	1.72	42.37**	-0.90	0.41	-0.85	0.88
l_f	3.75	43.04**	0.83	8.09**	-0.91	0.71	-1.95	4.08**

Panel C: Model Diagnostics

	Model 1		Model 2		Model 3		Model 4	
	Test	***p*-value**	**Test**	***p*-value**	**Test**	***p*-value**	**Test**	***p*-value**
R^2	0.07	--	0.01	--	0.78	--	0.69	--
F	14.09	0.00	0.10	0.75	63.66	0.00	165.9	0.00
Omnibus test	0.61	0.74	1.64	0.44	0.36	0.83	0.71	0.70
Q	12.34	0.49	5.36	0.97	11.66	0.56	14.83	0.32

Note: Significance at the 10 and 5% confidence level is indicated by * and ** respectively. All statistics are obtained from ML estimation of the VAR ECM representation with one lag in the original VAR in levels. Max(λ) and Trace statistics are likelihood ratio tests statistics for the following hypotheses:

Max(λ): H_0: rank(**Π**) = r *vs*. H_1: rank(**Π**) = $r + 1$

Trace: H_0: rank(**Π**) = r *vs*. H_1: rank(**Π**) > r

Critical values are obtained from Osterwald-Lenum (1992) for models 1 and 2 and from Pesaran *et al.* (2000) for models 3 and 4.

The max(λ) and the trace statistic for this model are presented in Panel A of Table 3 under the heading Model 1. The null hypothesis of the absence of a cointegrating relation between the endogenous variables **z** is rejected at the 95%

confidence level for both statistics. In addition, the null of the existence of just one cointegrating vector cannot be rejected at the 90% level for both statistics.

We thus estimate the ECM under the assumption of just one cointegrating equation. All signs of the estimates of the parameters in the CE were of the expected sign. In panel B, under the heading Model 1, we report Likelihood Ratio and Wald tests for their significance. All nonstationary variables were significant in the CE in at least one of the tests with the exception of capital services and wages.

Given the admittedly small sample size, it could prove useful for our simulation purposes to simplify the model by imposing some testable restrictions on the I(0) variables. For instance, the parameters related to the weather indicator are not significant at the 80% levels and for *CP* and *SA* we cannot reject the null gain hypothesis using a Wald test. We therefore drop Weather from the equation and impose null gain restrictions on the other exogenous I(0) variables by including them in differences. The resulting model provides estimates for the parameters of ∇CP and ∇SA which are not jointly significant: using a standard F test, the *p*-value of the test for the null of zero coefficient in the two variables is 14.5. We thus drop all exogenous variables and estimate Model 2, which consists only of the deterministic components and the cointegrating relationship. Note that, by construction, Model 2 cannot assess the effects of the MacSharry reforms since variables related to the reforms have been dropped from the original specification.

Cointegration rank tests for Model 2 are reported in Panel A of Table 3. As in Model 1, the tests unambiguously suggest the existence of one cointegration relation between the nonstationary variables. In addition, the LR and Wald exclusion tests presented in panel B indicate that the parameters for output price, price of land services, and price of capital services are all significant and of the expected sign (i.e. opposite to the sign for output price and of the same sign for prices of land and capital services). The estimated coefficients for input prices and wages are not significant, whereas the coefficients for family labor and TFP are again significant.

Panel C in Table 3 presents several model diagnostics tests for the output equation in the VAR representation. We cannot reject the null of absence of autocorrelation and normality in the residuals of the two models and for Model 1, the coefficients in the VAR representation are jointly significant. However, the adjusted R^2 is disappointingly low for both specifications.

In order to attain a more efficient multivariate analysis, given the relatively short time span of our data, we follow Pesaran et al. (2000) and assume that some of the integrated variables are exogenous. The set of exogenous I(1) variables was selected from the complete set of nonstationary variables based on the results, reported in Table 4, for the weak exogeneity test proposed by Engle (1984). Results are reported for the case of one lag. For all variables but the price of agricultural land services we cannot reject the null of weak exogeneity at the 10% level, although own price is clearly a bordeline case.[22] Therefore, we assume that all integrated variables are exogenous with the exception of the price of land services, which is the only agriculture-specific factor, and estimate equation (2). We also omit from the vector of exogenous I(1) variables those which were not significant in Model 1 in order to simplify our model specification. Thus, we arrive at Model 3 which refers to the following specification: The endogenous variables are q and r_V, the exogenous I(1) variables are p, p_M, l_f, and TFP, and the exogenous I(0) variables are *Weather*, *CP*, *SA*, and their lagged values.

Table 4: Engle's Weak Exogeneity Test.

DATA GENERATION PROCESS OF AGRICULTURAL OPUTPUT

Variable	Test	p -value
p	-1.6914	0.1032
p_M	0.6326	0.5328
r_V	2.2097	0.0365
r_K	-1.3768	0.1808
w	-0.0762	0.9399
Lf	0.3655	0.7178
TFP	-0.914	0.3694

Note: Results were obtained setting the number of lags to one. See the main text for a comment on the robustness of these results.

In columns 5 and 6 of Table 3, we present the cointegration rank tests for Model 3. As with models 1 and 2, there is clear evidence of the existence of one cointegrating vector. However, the LR and Wald tests for the significance of each

[22] Since the number of lags chosen potentially influence the output of the test, we carried it out using up to six lags. The results were similar to the one-lag specification and are not reported for reasons of brevity. For the output price variable, the p-value of the test narrowly fluctuated between 8.12 and 16.14, suggesting that it is weakly exogenous.

of the variables in the cointegrating vector give somewhat contradictory results, because we are unable to reject the null for any of the I(1) variables. We attribute this result to the small sample size and the very high correlations between the I(1) variables. For instance, if we dropped the variable with the highest *p*-value in the tests, p_M, then both p and *TFP* would became significant (not shown in the Tables).

Model diagnostics in Panel C refer to the estimation of the output ECM after imposing exclusion restrictions on the variables for which the parameters are not significant, that is, in the presence of comtemporaneus growth in prices, input prices and total factor productivity. Panel C in Table 3 shows that model 3 not only passes the diagnostics for normality and absence of autocorrelation, but also presents a significantly higher adjusted R^2 than models 1 and 2 so that it accounts for most of the observed variance in output growth. Not only are the slope coefficients jointly significant, as shown by the F-test reported in Table 3, but also each of the remaining coefficients is significantly different from cero.

Also, all stationary variables are significant and we reject the null of zero gain for *CP*. Thus, in our last model specification, Model 4, both *Weather* and *SA* appear in differences while *CP* is included in levels. Given the very high correlation between *SA* and *CP*, 0.997, and the fact that in this last specification its coefficient was not significant, we decided to drop *SA* in differences. Thus, caution should be exercised when interpreting the coefficients for *CP*, as they likely reflect the mixed effect of the compensatory payments and the set-aside requirements. To further increase the degrees of freedom in our analysis, we also drop from the VAR representation contemporaneous price growth as it is the only remaining regressor which is not significant. Finally, in the CE we keep output price, family labor, which is strongly significant both in the VAR and the CE, and total factor productivity together with the endogenous variables.

Results remain unchanged regarding the cointegration rank tests. With respect to exogenous variables in the CE, all variables are of the expected sign, although total factor productivity is not significant. As regarding the VAR representation, the comments for Model 3 apply as well, with the only difference that the adjusted R^2 is slightly lower, 69.39.

Both Model 3 and Model 4 are parsimonious representations that nevertheless can account for most of the variance in output growth. Thus, they are suitable for the counterfactual exercise that we report in the next section.

IV.3 Simulation Results

We cannot know with certainty the type of agricultural policies that Spain and other nations would have adopted after 1983 had Spain not entered the EU. As a result, we develop three counterfactual scenarios that provide differing perspectives on the impact of CAP policies on Spanish agriculture.

In the first scenario, Spain is assumed to continue the same agricultural policies in effect prior to 1983 when internal prices were either equal to world prices or above them by a nominal tariff rate. This time period was selected since Spanish policies had begun the process of converging to EU policies as early as 1983. To construct forecasts of Spanish tariffs under this scenario, we applied

ARIMA models to annual tariff data from 1963 to 1982 for each of 15 major Spanish agricultural products. Each forecasted-tariff rate was then aggregated into four broad product groups – cereals, vegetables, fruits and animal products -- using weights based on the importance of each product in the broader product group. We find that average Spanish tariffs stabilize at 22 percent for cereals, 3 percent for fruits, 0 percent for vegetables and 39 percent for animal products. Next, we developed world price indices for each of the four major product groups using U.S. price data for 56 products from 1973 to 2000 listed online by the U.S. Department of Agriculture.[23] Finally, we applied the forecasted tariffs to our international prices, and aggregated the resulting counterfactual internal price into one single Fisher price index.

In the second scenario, Spain is assumed to move unilaterally to free trade in agricultural products. This policy option is always available to nations so that its effects provide a useful contrast to those due to adoption of interventionist policies like those of the CAP. We again employ the scenario-one indices of world prices for outputs and inputs developed above, but Spanish tariff rates are, of course, set to zero.

[23] A second set of international prices were developed in order to check the adequacy of the price counterfactual. This second set of indices was constructed as follows. The price indices for the vegetables and fruit product groups were constructed using price data for 25 products from 1980 to 2000 listed in FAOSTAT, an online statistical database compiled by the Food and Agriculture Organization of the United Nations. The price indices for the cereals and animal products groups were constructed using price data from 1980 to 2000 for 13 products listed in the online database compiled by the International Monetary Fund. The resulting series were very similar to the U.S. data and consequently simulation results were unchanged to the first digit.

The third scenario takes the second scenario one step further by assuming that all nations liberalize their trade in agriculture. This scenario is of particular interest since pressure for free world trade in agriculture has grown more intense in recent years. New estimates of world agricultural prices are computed for this scenario since studies project that world agricultural prices will undergo change following the elimination of world trade barriers. We adjust world prices in each product group by estimates of the percentage increase in price due to liberalization as computed by Diao, Roe and Somwaru (2002), who project that world cereal prices will rise by 10 percent, fruit and vegetable prices will remain unchanged, and that animal products will rise 25 percent.

The first scenario also incorporates an evolution for input prices different from what actually occurred. In the absence of EU entry, it is plausible to assume that Spanish producers of feeds, fertilizers, agrochemicals, and seeds would not have faced the same degree of competition from EU producers. Thus, we develop a series for agricultural input prices that reflects the evolution of these prices after 1985 under the assumption that Spain had not entered the EU. Our first step was to obtain annual price data from 1970 to 2000 for feeds, fertilizer, agrochemicals, seeds, and energy from Eurostat's New Chronos online database. We next conducted a series of tests for structural breaks in the drifts for the individual price series. The tests were conducted by breaking each price series into two subperiods and then estimating ARIMA(0,1,2) models over each. A first set of tests for structural breaks in the years leading up to EU entry indicated that no breaks were present during this period. This indicates that pre-1986 input prices were not influenced by Spain's potential entry into the CAP and is consistent with the fact that Spanish input tariffs remained unchanged until entry. However, we rejected the null hypothesis of no structural break (at the 10% significance level) in 1986 for the three most important agricultural inputs: feeds, fertilizers and agrochemicals. This finding suggests that EU entry and access to European producers has reduced prices for these inputs. We were unable to reject the null hypothesis for seeds and energy prices. The inability to find a structural break for energy prices should not be surprising since EU entry has not facilitated access to crude energy supplies and input tariffs have not changed. For those price series that exhibited a structural break in 1986, we constructed the post-1986 evolution of the price series as if the structural break had not taken place. For seed and energy, the two input price series that had no structural break, we utilize the actual evolution of input prices after 1985. These five input price series were aggregated into a Fischer index of counterfactual input prices. Finally, all scenarios assume that the MacSharry reforms did not take place.

The accumulated growth rates of prices under each of the scenarios are listed in Table 5.

Table 5: Counterfactual Scenarios. Accumulated Growth: 1983-2000.

	p	p_M	*CP*	*SA*
Actual Growth	13.89	9.67	105.53	92.69
Scenario 1	4.26	19.62	0.00	0.00
Scenario 2	2.73	9.67	0.00	0.00
Scenario 3	5.27	9.67	0.00	0.00

Note: Figures are the accumulated growth from 1983 to 2000 in percentage terms for output price (p) and input prices (p_M) and the accumulated growth from 1992 to 2000 for compensatory payments (*CP*) and set-aside rates (*SA*). In the first scenario, government agricultural policies worldwide remain unchanged so that Spanish internal prices are above world prices by a nominal tariff rate consistent with previous Spanish trade policy. In the second scenario, Spain is assumed to move unilaterally to free trade in agricultural products and inputs. Finally, in the third scenario, all world trade barriers to agricultural products and inputs are removed and Spanish internal prices equal world prices in a liberalized world agricultural market.

As is to be expected, in all three scenarios prices grow well below the actual rates that prevailed under the CAP. However, since the prices of some important Spanish crops are projected to rise with free world trade, total liberalization (scenario 3) implies higher price growth for Spanish agriculture than under a policy of continued protectionism (scenario 1).

Our main results, the counterfactual simulations, are presented in Table 6. Using models 3 and 4 and the counterfactual price series, we have computed the percentage changes in output that would have occurred from 1983 to 2000 had the counterfactual prices under each scenario been in effect. These growth rates are listed in rows 3-5 and 7-9 in the first column of Table 6.

Table 6: Simulation Results. Accumulated Agricultural Output Growth: 1983-2000.

	Total (p, p_M, *CP and SA*)	Output Prices (p)	Input Prices (p_M)	1992 Reforms (*CP and SA*)
Actual Growth	4.2115	--	--	--
Model 3				
Predicted	4.4770	--	--	--
Scenario 1	5.7325	3.7507	4.2432	6.6927
Scenario 2	5.8383	3.6226	4.4770	6.6927
Scenario 3	6.0559	3.8402	4.4770	6.6927
Model 4				
Predicted	4.6719	--	--	--
Scenario 1	7.5498	3.9371	4.6719	8.2847
Scenario 2	7.4200	3.8072	4.6719	8.2847
Scenario 3	7.6208	4.0080	4.6719	8.2847

Note: See Section IV in main text for the definition of each model. Each column reports the predicted output growth in percentage terms assuming the counterfactual for the variables in the column header. See Table 5 for the definition of the different scenarios.

As a check of each model's ability to fit the data, we have also computed for each model the percentage change in output that would have occurred had actual prices been in effect. We refer to these as the 'predicted' growth rates and list them in rows 2 and 6 in the first column of the table. Contrasting these values with the actual growth that occurred over the period, 4.21 percent, we can conclude that the models perform well: the predicted growth rates for both models closely follow actual growth. In particular, after 17 years of simulation, Model 3 deviates from actual growth by only .27 (= 4.48 - 4.21) percentage points, whereas Model 4 also deviates only .46 (= 4.67 - 4.21) percentage points.

Turning now to the growth rates estimated for each scenario, it is immediately clear that in all cases, regardless of the model specification, Spanish agricultural output growth would have been higher if Spain had not implemented the CAP. For instance, under total liberalization (scenario 3), Model 3 indicates that output could have grown 6.06 percent, thus exceeding the predicted growth rate by 35.3 percent. Growth would have been even higher with Model 4, which projects a growth rate of 7.62 percent or 63.17 percent above the predicted rate. Moreover, even the least growth-enhancing scenario -- maintaining protectionist policies (scenario 1) -- would have raised growth by about 1.25 percentage points above the predicted rate. These results are consistent with those who argue that recent CAP policies have been effective in constraining EU agricultural output growth.

In order to identify the forces underlying the growth rates in each scenario, we re-compute them setting one variable's evolution according to the counterfactual for that scenario and the other two variable's evolution to their actual values. This process effectively removes the influence on output growth of a particular CAP policy as it operated through the variable. These growth rates are presented in columns 2 through 4. Thus, the first growth rate in column two, 3.75 percent, is below the predicted rate of 4.48 percent, which indicates that agricultural output growth in Spain would have been .73 percentage points lower if prices had not grown as they did under the CAP. This conclusion holds across all scenarios in both models; output growth would have been between 14.3 and 19.2 percent smaller than it actual was.

In column three, we present the evolution of output under the restriction that input prices would not have evolved as they did with the CAP in effect. Note that under scenarios 2 and 3, the counterfactual is equivalent to the actual series so that the re-estimated growth rates are equal to the predicted growth rates for each model. Although the two models, by construction, are not different, the re-estimated growth rate under Model 3 does indicate that if input prices had grown faster, then output would have increased less. However, this effect is very small (a difference of only .23 percentage points) and cannot explain why output would have increased much more under the counterfactual in spite of the price effects.

The answer to this question is in the fourth column of Table 6. Our simulations suggest that the 1992 reforms have had an important impact on agricultural output. Model 3 states that without them and even with the price increase for joining the EU, output growth would have been between (6.69-4.48)/4.48 x 100 = 49.8% higher than it actually was. The figure for Model 4 is 77.3%.

V. CONCLUDING REMARKS

In this paper, we analyzed the long-term relation between Spanish agricultural production and Spanish agricultural policies from 1970 to 2000, a period spanning Spain's implementation of the Common Agricultural Policy in 1986 and the application of the MacSharry Reforms in 1992. We did so by applying cointegration techniques to these data from which we derived two models, each capturing the long-term relation between prices and production. These models in turn permitted us to assess the agricultural production changes induced by the CAP by projecting the evolution of prices in the absence of Spain's entry into the EU in 1986. We therefore developed three plausible counterfactual scenarios for prices: in the first, Spain is assumed to maintain its pre-1986 agricultural policies, in the second it moves unilaterally to free trade, and in the third, all countries move to free trade in agricultural products. We find that Spanish agricultural output growth would have been higher under all three alternatives to the CAP for two main reasons: first, under each scenario, Spanish output is no longer constrained by the MacSharry reforms and, second, world liberalization of agricultural trade implies higher growth in prices for important Spanish crops.

These results are significant since they have a bearing on two key issues hotly debated within and outside of Europe: the price responsiveness of European agricultural supply and the effectiveness of the MacSharry policies in limiting agricultural production. Opponents of the CAP argue that European agricultural supply is price responsive and that the reduction of support prices would adequately stem European agricultural surpluses. Defenders of the CAP counter that dismantling the price support system, while doing nothing to stem agricultural surpluses, would certainly cause needless harm to farmers' income and hinder the EU's ability to attain other social and environmental benefits. Both groups concede that they are unsure of the implications of the set-asides and income transfers, an uncertainty also reflected in theoretical work.

Our results are supportive of both positions. Our models indicate that Spanish agricultural supply is responsive to changes in support prices, whereas our counterfactuals demonstrate that the MacSharry reforms have constrained Spanish agricultural production. Indeed, the counterfactuals indicate that the impact of the MacSharry production constraints has been substantial; had they not been in place, agricultural surpluses would have grown dramatically despite the freezing of CAP support prices. These results are encouraging for two reasons. First, they strengthen the position of those reformers both within and outside of Europe that argue for lower support price supports. Second, they indicate that recent EU reforms, which have in effect extended the MacSharry reforms, are appropriate measures for stemming agricultural surpluses.

APPENDIX: Data description

Spanish agricultural production data from 1970 to 2000 was obtained from EUROSTAT's New Chronos online database and is based on the Economic Accounts of the Agriculture following EUROSTAT's new methodology. Each annual work unit of family labor is the equivalent to the work of one relative working full-time in the farm during one year. The main source for the data is the Ministry of Agriculture yearly statistical books from 1979 to 2000 (MAPA, 1979-2000). Data prior to 1979 were obtained linking the series with data from *Encuesta de Población Activa*, *INE*, a rotating panel in which around 40,000 households are interviewed during 8 consecutive quarters.

Fischer price indexes were constructed for output and intermediate input prices. These data were obtained from EUROSTAT's New Chronos online database. For output prices, we took 24 main crops and animal products. For intermediate input prices, we collected data for seeds, agrochemicals, fertilizers, energy consumption, feeds, and other intermediate products. For 13 products, data were not yet available for the last two years, so we estimated these missing data points by applying the TRAMO-SEATS ML procedure, described in Gomez and Maravall (1994).

Land rental rates were computed as the ratio between average annual price of land to the long-run nominal interest rate. For the price of land, we used the figures provided by the Ministry of Agriculture's Survey of Land Prices on the price of Cultivated Land. We used the 2-year Treasury Bond interest rate from the Bank of Spain as the long-run nominal interest rate. Capital rental rates are 3-year moving average capital nominal rental rates computed using the long-term nominal interest rate and capital prices from Ball *et alia* (2003).

Gross hourly industry wage were obtained from the *Encuesta de Salarios, INE*, a quarterly wage survey available for the entire period. Total factor productivity is proxied by the five-year moving average of US agricultural productivity available online from Economic Research Service of the *USDA*. Rain and temperatures are computed from annual national averages from original monthly provincial data available at the *Instituto Nacional de Meteorología* (INE, 1960-2000). The data used are the annual deviations from historical averages from 1960 to 2000. Compensatory payments are available online at the Ministry of Agriculture web page while set-aside rates were obtained online from Eurostat.

The Spanish tariffs used in the simulation prior to 1983 were collected from García Alvárez-Coque (1986) and are equivalent producer's subsidies computed for the period between 1963 to 1982 for 15 major Spanish agricultural products.

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