Spanish agricultural exports competitiveness: The role of macroeconomic variables

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SUMMARY – This paper is one of the first attempts to analyse the effect of macroeconomic variables on the competitiveness of the Spanish agricultural sector. The objectives of this paper are twofold: first, it analyses the impact of exchange rate movements on agricultural prices and exports both in the short and long-run. Second we try do determine the relative importance of macroeconomic variables in agricultural prices instability. The methodological approach used is based on the cointegration procedure. In the empirical analysis, eight variables have been considered: exchange rate (ER), money supply (M), interest rate (R), the general price level (P), gross domestic product (Y), farm input prices (IP), farm output prices (OP) and total agricultural exports (AX). Quarterly data from 1978:1 to 1995:4 is used. To test for cointegration between these variables the Johansen's Maximum Likelihood approach is used. As money supply and the general price level are I(2), first the cointegration approach with I(2) system has been considered and neutrality cannot be rejected. This relationship has then been introduced in the I(1) system and variables have been defined in real terms. Then, in order to identify long-run relationships, restrictions on the individual cointegration vectors have been tested.

Key words: Competitiveness, macroeconomic variables, agriculture, cointegration, identification, short-run, long-run.

RESUME – *"La compétitivité à l'exportation du secteur agricole espagnol : Le rôle des variables macro-économiques". Le présent papier constitue l'un des premiers essais dans l'analyse de l'effet des variables macro-économiques sur la compétitivité du secteur agricole espagnol. L'objectif est double : d'une part analyser l'impact des variations du taux de change sur les prix agricoles et sur les exportations à court et à long terme, et, d'autre part, on essaye de déterminer l'importance relative des variables macro-économiques dans l'instabilité des prix agricoles. L'approche méthodologique utilisée est basée sur les techniques de cointégration. Dans l'analyse empirique, huit variables ont été considérées : taux de change (ER), offre de monnaie (M), taux d'intérêt (R), indice général des prix (P), produit intérieur brut (Y), prix des inputs agricoles (IP), prix des outputs agricoles (OP) et exportations agricoles totales (AX). Des données trimestrielles couvrant la période 1978/1 à 1995/4 ont été utilisées. L'approche "Maximum de Vraisemblance de Johansen" a été utilisée pour tester la cointégration entre les différentes variables. Etant donné que l'offre de monnaie et l'indice général des prix sont I(2), la technique de cointégration avec un système I(2) a été utilisée et la neutralité n'a pu être rejetée. Cette relation a été introduite dans un système I(1) et les variables ont été définies en termes réels. Dans le but d'identifier les relations à court et à long terme, des restrictions sur les vecteurs individuels de cointégration ont été testées.*

Mots-clés : *Compétitivité, variables macro-économiques, agriculture, cointégration, court terme, long terme, identification.*

Introduction

Changes in the macroeconomy have become increasingly significant within the agrofood sector as agriculture has become more capitalized and more dependent on international markets, then being more vulnerable to variations in interest rates, exchange rates and international growth rates.

During the 1980s the linkages between macroeconomic policies and agricultural trade have been an important research issue of agricultural economists. Since the early work of Schuh (1974), who first pointed out the importance of these variables in determining the exchange rate and the competitiveness of the agricultural sector, a number of papers have considered these relationships. The literature can be classified into three main groups taking into account the type of linkage between macroeconomics and agriculture considered: (i) one group of papers deal with the responsiveness of agricultural commodity prices and total agricultural exports to exchange rate movements (Chambers and Just, 1982, 1986; Barnett *et al.*, 1983; Andrews and Rausser, 1986; Orden, 1986; Devadoss *et* *al.*, 1987; Staoulis and Rausser, 1988; Orden and Fackler, 1989; Taylor and Spriggs, 1989; Carter *et al.*, 1990; Adamowicz *et al.*, 1991); (ii) a second group study the effects of monetary supply shocks on the agricultural and non agricultural prices (Bordo, 1980; Tweeten; 1980; Chambers and Just, 1982; Starleaf, 1982, 1984; Starleaf *et al.*, 1985; Bessler and Babula, 1987; Devadoss and Meyers, 1987; Devadoss *et al.*, 1987; Sephton, 1989; Taylor and Spriggs, 1989; Robertson and Orden, 1990; Adamowicz *et al.*, 1991; Denbaly and Torgerson, 1991; Larue, 1991; Larue and Babula, 1994; Jeffrey and Lastrapes, 1996); and (iii) a third group concentrate on the relationships between interest rates and commodity markets (Shei, 1978; Schuh *et al.*, 1980; Freebairn *et al.*, 1982; Chambers, 1983, 1984; Orden, 1986; Devadoss *et al.*, 1987; Orden and Fackler, 1989).

This paper is one of the first attempts to analyse the effect of macroeconomic variables on the competitiveness of the Spanish agricultural sector. The paper covers the three relationships mentioned before. First, it analyses the impact of exchange rate movements on agricultural prices and exports both in the short and long-run. Second, the impact of interest rates on the agricultural sector is considered. Finally, the relative importance of macroeconomic variables on agricultural prices instability is studied. In other words, to test whether agricultural prices respond faster or slower than input prices to a shock in monetary policy. The methodological approach used in this paper is based on the cointegration analysis which has been developed in the last decade and which allow us to examine long-run equilibrium relationships as well as short-run dynamics.

To achieve the mentioned objective, the paper is organized as follows. First, there is a section where we will examine the three basic theoretical linkages that will be considered in the empirical analysis. Second, there is a section where data used in this study is described and the univariate properties of series are investigated. Third, the long-run equilibrium relationships are analysed in another section. In the last section the short-run dynamics is considered. Finally, some conclusions are outlined.

Theoretical relationships between macroeconomics and the agrofood sector

Exchange rates effects

The relationship between exchange rates and agricultural trade (exports and prices) has been the subject of somewhat controversial literature. In the 1980s a series of theoretical and empirical models were developed to investigate the impact of exchange rates movements on agricultural foreign trade. Under floating exchange rates, a currency movement (appreciation or depreciation) leads to short-run adjustments in prices, output and trade volume considering a perfectly competitive market. Generally, market shares and export volume shift away from the country whose currency is appreciating and towards the country whose currency is depreciating. However, in the long-run, as the monetarists point out, relative prices will be unchanged by the movement of exchange rate because the *Law of One Price* will hold. Models can broadly divide into two groups: those in which the exchange rate is considered a purely exogenous variable; and those in which it is endogenously determined. In the early empirical work dealing with the effects of changes in the exchange rate on the export volume and domestic prices, single market (partial equilibrium) models were specified and estimated. Normally, the exchange rate was the only macroeconomic variable included in the model and was introduced as purely exogenous.

Under some restrictive assumptions (one market, one good and the exchange rate as a predetermined variable), the elasticity of domestic price with respect to the exchange rate lies between –1 and 0[1]. In other words domestic price changes will tend to be smaller or equal than the percentage of depreciation (or appreciation) of the exchange rate. On the other hand, the percentage change on export volume may be greater, lower or equal than the percentage change in price depending on if excess supply is elastic, inelastic or the unity, respectively.

[1]In a single-market partial equilibrium analysis, the effect of a change in the exchange rate on the equilibrium domestic price is:

$$-1 \leq \lambda_i = \frac{-\eta^*_{ii}}{\eta^*_{ii} - \varepsilon_{ii}} < 0$$

where η^*_{ii} is the own-price elasticity of the foreign excess demand and ε_{ii} is the own-price elasticity of home country excess supply (see Chambers and Just, 1979).

However, Orden (1986) shows, theoretically, that if the exchange rate and national income are included as endogenous variables, this elasticity will not be restricted between –1 and 0. Chambers and Just (1986) stress more general models and show that, in theory, the admissible exchange rate elasticity of agricultural prices may be even less restrictive if interest rates are also endogenized.

Obviously, this approach is more interesting as many macroeconomic variables are highly endogenous. Let us briefly comment this point. The traditional keynesian view of how a monetary tightening[2] is transmitted to the real economy is $M\downarrow \Rightarrow r\uparrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$, where $M\downarrow$ indicates that a contractionary monetary policy leads to a rise in the interest rate ($r\uparrow$), thereby causing a decline in investment ($I\downarrow$), causing a decline in aggregate demand and a fall in output ($Y\downarrow$). On the other hand, exchange rate movements are an important element in the transmission of monetary policy to the economy. When domestic real interest rates rise, domestic currency deposits become more attractive than deposits in foreign currencies leading to an appreciation of the exchange rate. As a consequence, the higher value of the domestic currency makes domestic goods more expensive than foreign goods (losing competitiveness), and this causes aggregate output to fall ($M\downarrow \Rightarrow r\uparrow \Rightarrow E\uparrow \Rightarrow EXP\downarrow \Rightarrow Y\downarrow$).

Interest rates effects

Interest rates can influence the farm sector in two ways: (i) cost effects; and (ii) stock effects.

As a consequence of a monetary contraction, the cost of production will rise due to the increase in interest rates and, in the absence of government subsidies, agricultural production will fall. On the demand side, the rise in interest rates will cause a short-run contractionary effect on income, reducing consumption, and this phenomenon will produce a downward pressure on domestic prices. In the same way, higher interest rates will affect farmers in the sense that borrowed money will be more expensive and storage costs will raise. Increasing of storage costs will cause a decrease inventories by releasing stocks onto the market (supply rise) and this will exert a downward pressure on domestic prices and agricultural producers will lose competitiveness from both the price and cost effects. The importance of interest rates effects on the farm sector has been suggested by many empirical studies (Shei, 1978; Schuh *et al.*, 1980; Freebairn *et al.*, 1982; Chambers, 1983, 1984).

Effect of inflation and money supply

The third aspect that has received much attention in the literature of agricultural economics is the effect of changes in monetary supply on relative agricultural prices (Bordo, 1980; Chambers and Just, 1982; Chambers, 1984; Orden and Fackler, 1989; Roberston and Orden, 1990; Larue, 1991; Larue and Babula, 1994). This question originate some controversy centred around whether or not monetary phenomena can have real effects and if inflation is neutral with respect to agricultural prices.

According to the monetarist approach, an increase in the money supply raise the general price level in the same proportion, then money is neutral in the long-run; that is, money does not affect the relative price level in the long-run. However, in the short-run prices do not adjust at the same velocity. The adjustment speed of different prices to a money supply shock depends on the macroeconomic theory considered (classical macroeconomic theory or Keynesian theory).

The basic assumptions of new classical macroeconomic theory assume that money is neutral. Likewise, real interest rate is not affected by monetary policy and nominal exchange rate will respond to changes in the expected inflation rate quickly, according to Fisher equation. The classical approach developed above is relevant for a world of relatively homogenous commodities sold in competitive Walrasian "auction market". In such markets, prices are assumed to be flexible enough to response to changing demand and supply conditions.

In contrast, the Neo-Keynesian paradigm considers the existence in the economy of different degrees of price flexibility and, then, differentiates two sectors: a fixed-price sector where price adjust

[2]A expansionary monetary policy work in opposite direction $M\uparrow \Rightarrow r\downarrow \Rightarrow I\uparrow \Rightarrow Y\uparrow$.

sluggishly to demand shocks; and a flexible price sector responding quickly to any shock. The existence of both sectors implies that any policy shock (money supply) leads to a change in the relative prices between both sectors which will affect real output, interest rates and exchange rates, and which may overshoot[3] the long-run equilibrium values in response to policy shocks (Dornbusch, 1976; Frankel, 1986).

Whether monetary policy has significant effects on agricultural relative prices has been a crucial question in many recent discussions of macroeconomic impacts on agriculture. This discussion has its origin when considering the agricultural sector as a competitive one in which prices are more flexible than those in non agricultural (fixed-prices) sectors. Under this hypothesis, it has been argued that commodity prices respond quicker than manufactured goods prices to changes in the money supply. This hypothesis suggest that farmers should benefit from an unanticipated increase in the rate of inflation and this will make the agricultural sector more competitiveness, at least in the short-run. A clear example of this result is in Bordo (1980) and Frankel (1986). However, the opposite hypothesis has been supported by Tweeten (1980) who concluded that agricultural prices are less responsive than manufacturers prices to an increase in money supply. This author argues that domestic inflation raises costs while prices remains unchanged then framers are subject to "cost-price-squeeze". Finally, Larue and Babula (1994) found empirical evidence of both approaches in USA and Canada depending on the time period considered.

Data and methodological approach

To carry out the empirical analysis of the linkages between the agricultural sector and macroeconomics two blocks of variables have been considered. The first one is the macroeconomic block and contains the more relevant macroeconomic variables: exchange rate (ER)[4]; money supply (M); interest rate (R), the general price level (P) and gross domestic product (Y). The second one is the agricultural block which includes the following variables: farm input prices (IP), farm output prices (OP) and total agricultural exports (AX). Quarterly data from 1978:1 to 1995:4 is used. These variables were chosen because it was felt that they would capture the most important relationships between both sectors[5] and, on the other hand as the sample period is limited, we have attempted to use as few variables as possible. All variables are in logarithms, except for the interest rate which is divided by one hundred to transform them into the same order of magnitude as the differenced logarithmic variables. Finally all variables have been deseasonalized.

A Vector Autoregressive Model (VAR) model containing the eight variables has been specified. The VAR model approach relatively new in studying the macroeconomic-agricultural linkages but it has gained a lot of popularity in the last years. In general terms, the VAR approach has been used to analyse the short-run dynamics of a shock in one variable on the other variables through the impulse response functions and the decompositions of the forecast error variances. One of the advantages of this methodology is that all variables are considered as endogenous and no zero/one restrictions are imposed on the variables in the system. The main criticism is that it is a purely a theoretical approach although what it is called structural VAR models have overcome this criticism.

Most economic analysis and empirical work, which have attempted to study the relationships between macroeconomics and agricultural sector, have focussed on short-run analyses. However, recent advances in time series modelling have provided us the ability to test for the presence of long-run equilibrium relationships between sets of variables by using cointegration tests. The concept of cointegration was popularised by Engle and Granger (1987) and was the starting point of a lot of

[3]Overshooting is a more than proportionate short- run response of nominal variables (like price, exchange rates) to a change in money growth.
[4]Money supply (ALP has been used), exchange rate (nominal effective exchange rate has been used), and interest rate (of three month has been used) have been obtained from the "Boletin de Estadística del Banco de España". Domestic national product (Y) and general price level (P) have been obtained from the "Boletines estadísticos del Instituto Nacional de Estadística (INE)". Input prices (IP) and output prices (OP) index have been obtained from "Boletines Mensuales de Estadística Agraria" of Agriculture, Fisheries and Food Ministry (MAPA) and agricultural exports (AX) has been obtained from "Dirección General de Aduanas".
[5]The foreign direct investment was not included in the analysis, as the available data only covers the period 1987-1995 and because, following the recommendations from the Ministry of Economy, in the first two or three years data are not very reliable.

literature of both theoretical and empirical research which has demonstrated the relevance for modelling the long-run behaviour. More recently, Johansen (1988) and Johansen and Juselius (1990, 1992, 1994) develop an interesting methodology which jointly consider both the short and the long-run. One of the most important advantages of this approach is the identification of structural relationships among a set of variables in the long-run taking into account the short and long-run adjustments. This is the methodological approach we are going to consider in this paper.

The first step in this approach consists of checking the univariate properties of data. Apart from the visual inspection of the graphs of the series and correlograms, unit root tests developed by Dickey and Fuller (1979, 1981) (DFA) and Kwiatkowski *et al.* (1992) (KPSS) have been applied. All variables are non stationary and most of them are I(1)[6]. The general price level (P) and money supply (M) are I(2). This complicates considerably the analysis of the cointegrating relationships. In the following section, the specification, estimation and identification of long-run relationships are analysed. Special attention will be paid to the analysis of cointegration with I(2) variables as it is somewhat more complicated that the analysis with I(1) variables and very limited applications have been found in the literature.

Long-run relationships

Cointegration with I(2) variables

As two variables in the system are I(2) we will start with some methodological issues concerning cointegration with this kind of variables. A complete description of this methodology can be found in Johansen (1995). When a VAR model includes I(2) variables a different type of cointegration can appear. I(2) variables may be cointegrated forming a I(1) space, and this relationship may be again cointegrated with other I(1) variables to form stationary relationships. This phenomenon is called multicointegration.

The starting point of Johansen's (1995) is a VAR model:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + \dots + \Pi_k z_{t-k} + \varepsilon_t \qquad [1]$$

where Π_i, i = 1, 2...,k, is an (pxp) matrix of parameters and ε_t is a vector of residuals which is assumed to be i.i.d Gaussian process ~iid(0,Σ) that is:

$$E(\varepsilon_t) = 0 \qquad \text{for all } t$$

$$E[\varepsilon_t \varepsilon_s] = \begin{cases} 0 & \text{if } t \neq s \\ \Sigma & \text{if } t = s \end{cases}$$

since Σ is a covariance (p×p) matrix positive definite.

This type of model is an unrestricted VAR expressed as a reduced form from which each variable Z_t is regressed on lagged values of itself and on lagged values of each of the (p-1) remaining variables. Defining Δ = 1-L, where L is the lag operator, equation [1] can be reformulated into a Vector Error Correction Model (VECM) in second difference in the following way:

$$\Delta^2 Z_t = \Gamma \Delta Z_{t-1} + \Pi Z_{t-2} + \sum_{i=1}^{k-2} \Psi_i \Delta^2 Z_{t-i} + \varepsilon_t \qquad [2]$$

In the I(2) system Z_t is said to be cointegrated if the following two reduced rank conditions are satisfied (Johansen, 1995):

$$H_{r,s} : \Pi = \alpha\beta' \text{ of rank } r<p$$

and [3]

$$H_{r,s} : \alpha'_{\perp}\Gamma\beta_{\perp} = \varphi\eta' \text{ of rank } s<p-r$$

[6]Results are not shown due to space limitations. They are available upon request.

where α and β are matrices of dimension p×r; $\alpha_\perp$ and $\beta_\perp$ are p×(p-r) matrices which are orthogonal matrices to α and β such that $\alpha'_\perp\alpha = 0$ and $\beta'_\perp\beta = 0$, and φ and η are (p-r)×s matrices of rank s (s<p-r).

In order to characterize the stochastic trends driving the I(2) system and the cointegration relations, it is necessary to define parameters describing the I(0), I(1) and I(2) relationships among the variables. The associated dimension of each sub-system is given by: r representing the number of cointegration vectors; s the number of I(1) components in the model; and s_1 = p-r-s, the number of I(2) components in the system. The values of r, s and s_1 are the so-called integration indices of the VAR (Paruolo, 1996). Following Johansen (1995), $\beta_\perp$ and $\alpha_\perp$ can be decomposed into:

$$\beta_\perp^1 = \beta_\perp\eta \;;\; \beta_\perp^2 = \overline{\beta}_\perp\eta_\perp \;;\; \alpha_\perp^1 = \alpha_\perp\varphi;\; \alpha_\perp^2 = \overline{\alpha}_\perp\varphi_\perp$$

where a line above the matrix β indicates that $\overline{\beta} = \beta(\beta'\beta)^{-1}$ and $\beta'\overline{\beta} = I$ (the same for α).

Considering these definitions, the different cointegration possibilities in the I(2) system are given by (Haldrup, 1998):

r I(0)-relations	$\begin{cases} s_1 = p - r - s \\ \text{and} \\ r - s_1 \end{cases}$	$\beta'Z_t + \delta(\beta_\perp^2)'\Delta Z_t$ $\delta'_\perp\beta'Z_t$
s I(1)-relations	$(\beta_\perp^1)'Z_t$	
s_1 I(2)-relations	$(\beta_\perp^2)'Z_t$	

where $\delta = \overline{\alpha}'\Gamma\overline{\beta}_2$ is a matrix of dimension r×s_1 and $\delta_\perp$ is a r×(r-s_1) matrix orthogonal to δ.

Then, it is easy to see that the combinations (β, $\beta_\perp^1$)ΔZ_t are stationary vectors and, then the estimates of β and $\beta_\perp^1$ define the relationships which reduce the integration order from 2 to 1 [CI(2,1)]. Thus, the cointegration parameters of interest in the I(2) system are given by β, $\beta_\perp^1$ and δ. On the other hand, the (p-r-s) combinations $\beta'Z_t + \delta\beta_\perp^{2\prime}\Delta Z_t$ are stationary and capture what it is called the multicointegration relations.

In addition, $\beta_\perp^2$ gives the weights with which the I(2) components affect the variables of the system (Juselius, 1994), whereas $\alpha_\perp^2$ determines the common stochastic I(2) trends.

Johansen (1995) proposes a two-step procedure to determine the dimensions of r, s and s_1. In the first step, Johansen (1995) proceeds as in the I(1) case to obtain the number of cointegration vectors (r) in the system and the estimation of α and β (the statistic used is called Q_r). The only difference relies on how residuals and their second moment matrices are calculated (see Johansen, 1995, for further details). Assuming that the parameters $(\hat{r},\hat{\alpha},\hat{\beta})$ are known, the second step involves of determining the rank of $\alpha'_\perp\Gamma\beta_\perp$ (s = 0,1,...,p-r-1)[7]. In this case, the statistic used is called $Q_{r,s}$.

The cointegration ranks, r and s, can also be jointly determined by applying the statistic $S_{r,s}$. This statistic involves testing the hypothesis $H_{r,s}$ against the unrestricted VAR model (r = p) and is given by:

$$S_{r,s} = Q_{r,s} + Q_r \qquad [4]$$

In this case, the joint hypothesis $H_{r,s}$ will be rejected if $H_{i,j}$ is rejected for (i<r) and for i = r and j≤s.

[7]Note that estimates are available for each value of r.

As a consequence, the chosen values (r, s) will correspond to the first hypothesis $H_{r,s}$ that is not rejected.

Once the methodological issues on cointegration with I(2) variables has been described, we are going to apply this methodology. The first step in the estimation of the I(2) system is to determine the lag length of the unrestricted VAR model containing only the macroeconomic variables. The Tiao-Box (1981) likelihood ratio test statistic[8] showed that the model lag length was three periods. A deterministic trend has been introduced in the long-run component[9]. Results from the tests statistics Q_r, $Q_{(s,r)}$ and $S_{r,s}$ are shown in Tables 1 and 2.

Table 1. Determination of multicointegration rank (r and s) in I(2) system

p-r	r	$Q_{(s,r)}$ [†, ††]				Q_r	CV(10%)
4	0	117.86 s = 0	41.32 s = 1	10.82 s = 2	4.20 s = 3	76.99	59.02
3	1		86.32 s = 0	10.82 s = 1	4.36 s = 2	40.00	39.26
2	2			68.31 s = 0	4.37 s = 1	18.43	22.98
1	3				4.97 s = 0	6.65	10.63
p-r-s = s_1		4	3	2	1		
CV(10%)		59.02	39.26	22.98	10.63		

[†]The critical values (10% significance level) are obtained from table 4 in Jorgensen *et al.* (1996).
[††]r is the number of cointegration vector, s is the number of I(1) components and (p-r-s = s_1) the number of I(2) components.

The right hand of Table 1 gives the results. As it can be observed, we fail to reject that r = 2. Once, r is determined, the $Q_{r,s}$ statistic is used to get the value of s. Results are included in the central part of Table 1. At the 10% of significance level, we fail to reject the null hypothesis that s = 2. However, the results of $S_{s,r}$ indicate that at the 10% significance level the hypothesis that (r,s) = (0,s), is rejected for all s. On the other hand, we fail to reject the hypothesis that (r,s) = (1,1) since the set value of $S_{1,1}$ = 50,81 which is under the critical value of 64,23. In this case, r = 1 and s = 1 and then, s_1 = p-r-s = 2 I(2) stochastic trends in the data. Given that Y and R are I(1) variables, while M and P are I(2), the s_1 = 2 indicate that (M-P) is a I(2) component. In other words, the money supply and the general price level would not be cointegrated, which is not economically reasonable[10]. Then, the next step is to test the hypothesis (r,s) = (2,1) which cannot be rejected at 10% significance level. This indicates that we have two cointegration vectors, one (s = 1) I(1) component and one (s_1 = p-r-s = 1) I(2) component and, in consequence we can say that money supply and the general price level are cointegrated.

Once, we have carried out the cointegration test, the next step will be to identify the two cointegrating vectors we have obtained. The purpose is to test if the price homogeneity or money neutrality holds. If we fail to reject this hypothesis, we will be able to model an I(1) system including all the macroeconomic and agricultural variables which will simplify the methodology used as it is more standard.

[8]Batteries of tests have been used to check for the presence of serial correlation and normality with different lag structures. The multivariate and univariate tests of serial correlation and normality provided no evidence of misspecification when 3 lags are considered.
[9]The I(2) analysis of a model allowing for a trend term in the long-run is given by Jorgensen *et al.* (1996) and Paruolo (1994). For more details of this approach see Jorgensen *et al.* (1996).
[10]Juselius (1994) suggest that the cointegration results should be economically reasonable. Thus she recommended the use of any prior economic insight on this matter such that the choice of r is consistent with the statistical information and gives an economically coherent interpretation of the number of long-run relations and common trends. The problem of small samples has also been considered as Reimers (1992) suggest that in such situations the Johansen procedure over-rejects when the null is true.

To identify the cointegration vectors, we followed the method proposed by Johansen and Juselius (1994) for the I(1) case. Jorgensen *et al.* (1996) show that in the I(2) case the same procedure applied, but, with a different interpretation as $\beta' Z_t$ is I(1)[11]. Let us briefly describe the method.

Table 2. Results from the rank test $S_{r,s}$ in the I(2) system.

p-r	r	$S_{r,s}$ †,††			
4	0	194.85 *(132.02)* s = 0	118.30 *(107.91)* s = 1	88.81 *(87.90)* s = 3	81.18 *(71.33)* s = 4
3	1		126.31 *(82.29)* s = 0	50.81 *(64.23)* s = 1	44.36 *(49.69)* s = 2
2	2			86.74 *(44.52)* s = 0	22.80 *(31.61)* s = 1
1	3				11.61 *(17.59)* s = 0
p-r-s = s_1		4	3	2	1

†Values in parenthesis are the critical values at the 10% significance level obtained from table 4 in Jorgensen *et al.* (1996).
††s is the number of I(1) components and (p-r-s = s_1) the number of I(2) components.

In order to identify long-run relationships, some restrictions has to be imposed on individual cointegration vectors ($\beta_i' Z_t$; $i = 1, \dots, r$) though the definition of restriction matrices H_i and R_i the hypothesis about β adopt the following general expression:

$$H_{0,\beta} : \beta = (H_1\varphi_1, H_2\varphi_2, \dots, H_r\varphi_r) \quad [5]$$

where the matrices H_i is a ($p\times s_i$) matrix, s_i is the number of unrestricted parameters in β_i and φ_i is ($s_i\times 1$) vector of parameters to estimates in the i^{th} cointegration relation. This hypothesis require the specification of ($p x p_i$) matrix R_i orthogonal to H_i such that $R_i'\beta_i = 0$ (p_i is the number of restriction imposed in vector β_i, that is, $s_i + p_i = p$).

Following Johansen (1992) the first vector in the cointegration space β_1 is identified if the following rank condition satisfied:

$$\text{Rank } (R_1'\beta_1, R_1'\beta_2, \dots, R_1'\beta_r) = \text{Rank } (R_1'H_1\varphi_1, \dots, R_1'H_r\varphi_r) = r - 1 \quad [6]$$

which means that is not possible to find a linear combination of vectors (β_2, ..., β_r) similar to vector β_1. This condition can be generalized as follow:

$$\text{Rank } \left(R_i H_{i1}, \dots, R_i H_{im}\right) \geq m \quad [7]$$

$i = 1, 2, \dots, r$; $m = 1, 2, \dots, r-1$ and $1 \leq i_1 \leq \dots \leq i_m \leq r$ (i excluded)

If condition [7] holds for a particular i, it means that restrictions in R_i identify vector i. In our case, in which r = 2, the rank condition is reduced to:

$$\begin{cases} \text{rank } (R_1'H_2) \geq 1 \\ \text{rank } (R_2'H_1) \geq 1 \end{cases} \quad [8]$$

[11]For more details see Theorem 5-2 (pp. 14-15) in Jorgensen *et al.* (1996).

To get an exact identification of cointegration space, not only [8] has to be satisfied, but, also that the number of restriction imposed (p_i) in each cointegration vector has to be exactly equal to r-1 ($p_i =$ r-1), or equivalent to $\sum_1^r p_i = r(r-1)$. The cointegration space is over-identified if more than r-1 restrictions are imposed on each cointegration vector, that is p_i>r-1, or $\sum_{i=1}^r p_i > r(r-1)$. In that case, it is always possible to test if restrictions are empirically identified through a likelihood ratio (LR) test statistic with $v = \sum_1^r (p - r + 1 - s_i)$ degrees of freedom.

In our case, the restrictions placed on the cointegration relations to test the long-run price homogeneity are formulated considering the following models:

$$\begin{aligned} \beta_1' Z_t &: M_t - P_t + \beta_{13} Y_t + \beta_{14} R_t = \mu_1 \approx I(1) \\ \beta_2' Z_t &: \beta_{23} Y_t + \beta_{24} R_t + \beta_{25} t = \mu_2 \approx I(0) \end{aligned} \qquad [9]$$

Thus, under these restrictions, the matrices H_1, H_2 are

$$H_1 = \begin{pmatrix} 1 & 0 & 0 \\ -1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & 0 \end{pmatrix}; H_2 = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \qquad [10]$$

The first is identified as a real money demand function while the second cointegration vector is identified as an output supply (IS equation) in which the real output variable is a function of interest rates and a linear trend. The estimated parameters of β under the restriction hypothesis in [10] are shown in part A of Table 3. The cointegration space represented by $\beta = (\beta_1, \beta_2)$ is generically identified since the necessary rank condition are satisfied[12]. However, in this case over-identification exists, since the number of restrictions p_i are higher than r-1. Then, we have to test if the model is empirically identified. As mentioned, in the case of over-identification, the test of whether restrictions are valid is a LR test distributed as $\chi^2(2)$. The LR value is 3.84, which is not significant at the 5% level. Therefore, the long-run homogeneity hypothesis is strongly supported by the data.

Table 3. Long-run identification in the I(2) system

Variables	A		B	
	β_1	β_2	β_1	β_2
M_t	1.000	0.000	1.000	0.000
P_t	–1.000	0.000	–1.000	0.000
Y_t	–1.290	1.000	–1.000	1.000
R_t	0.499	–7.609	0.668	2.537
T	0.000	–0.073	0.000	0.017
Q(f)	Q(2)		Q(3)	
	3.84		4.28	

By normalizing the first vector by M, the coefficients suggest both price and income homogeneity. Income homogeneity can be further investigated by setting $\beta_{13} = -1$. Following the same procedure as in the former case we have defined the appropriate H_i matrices. Results are shown in part B of Table 3. Also, in this case the rank conditions are satisfied. The value of the LR test is 4.28, which is well

12 $\begin{cases} \text{rank } (R_1' H_2) = 2 \\ \text{rank } (R_2' H_1) = 1 \end{cases}$

under the critical value at the 5% significance level. This indicates that both prices and income homogeneity are supported by the data. However, this hypothesis will be further investigated in the next section.

Cointegration in the I(1) system

Having accepted the long-run price homogeneity from the I(2) cointegration analysis, in this section we will carry out the cointegration analysis in real terms. The data base have been changed as follows: first the variables (M, IP and OP) have been deflated by the general price level (P); second, the general price level has be expressed in first differences (ΔP) and, finally the total agricultural exports has be deflated by an index of export prices. All the other variables remain the same[13]. Unit root tests indicate that all variables were I(1). A VAR model of I(1) variables has been specified. The Tiao and Box (1981) likelihood ratio test indicated that the optimum lag length was 3 periods. However, the results from multivariate normality and autocorrelation tests suggested that model was misspecified. Johansen and Juselius (1994) and Juselius (1994) suggest that non-normality and autocorrelation are due to wrong specification of the deterministic components in the model. These authors argue that the introduction of dummy variables can avoid the non-normality problem and therefore improve its stochastic properties. In this application, two dummy variables (D1 and D2) have been introduced in the short-run.[14] Tests on residuals indicate that, in this case, model was correctly specified.

Once the VAR model has been specified, the classical Johansen and Juselius (1990, 1992) procedure has been used. First, tests for cointegration has been used. However, as some dummy variables have been introduced, critical values are no longer valid. The Johansen and Nielsen (1993) procedure has been used to simulate the new critical values of the trace statistic. Table 4 shows the results. The null hypothesis that there are four cointegration vectors cannot be rejected at 5% level of significance.

Table 4. Results from Johansen Multivariate Cointegration tests

Eigen-values	H0: r =	p-r	Trace	λ-max	VC(90%)[†] Trace	VC(95%)[††] Trace
0.710	0	8	273.47	85.51	159.48	153.39
0.564	1	7	187.95	57.38	126.58	121.46
0.537	2	6	130.56	53.15	97.18	92.75
0.424	3	5	77.42	38.03	71.18	67.04
0.234	4	4	39.34	18.47	49.65	45.35
0.169	5	3	20.87	12.81	32.60	27.09
0.100	6	2	8.06	7.27	17.85	12.49
0.014	7	1	0.79	0.79	7.52	–

[†]Critical values are obtained from Osterwald-Lenum (1992).
[††]Critical values are obtained by simulation in Disco program of Johansen and Nielsen (1993).

The estimated parameters of α and β matrices obtained from applying the Johansen technique for r = 4 are presented in Table 5, where β is presented in normalised[15] form. The interpretation of the

[13]Now: RM is real money demand; ΔP is the first difference of the general price level; R is the interest rate; Y is the real gross domestic product; ROP the relative output price index; RIP is the relative input prices index; ER is the exchange rates index; and RAX is the real agricultural exports.
[14]D1 takes the value 1 over period 1978:1-1980:4 and zero in other case, while D2 takes the value 1 over the period 1992:1-1992:4. In the first case, we have taken into account the structural change that took place in the agricultural sector which faced higher production costs. In the second, we account for the devaluation process that took place those years.
[15]The normalisation chosen of the β matrix was arbitrary. Notice that any normalisation of β lead to different values in the α matrix.

unrestricted cointegration space is far from straightforward when there is more then one cointegration vector (Juselius, 1994). Moreover, Johansen and Juselius (1994) suggest that only sometimes the unrestricted cointegration vectors, surprisingly, can give a direct interpretation in terms of hypothetical relations. In the following lines we will give an economic interpretation of the cointegration vectors using the identification procedure mentioned in the I(2) system. As a first step, let as considered the unrestricted normalised vectors show in Table 5.

Table 5. Estimation of β and α matrices with r = 4

	β			
	Vector_1	Vector_2	Vector_3	Vector_4
RM	1.000	–0.217	–5.834	36.036
Y	–1.171	0.337	11.162	–56.544
ΔP	24.089	1.000	15.242	–26.797
ER	–1.368	0.006	–0.471	7.236
R	0.178	0.013	–1.312	1.581
RIP	–0.103	0.561	–0.492	3.366
ROP	1.455	–0.091	1.000	0.871
RAX	–1.205	0.035	0.483	1.000
Constant	11.283	–0.713	–29.629	83.385
	α			
RM	–0.041	0.485	–0.019	–0.010
Y	–0.012	–0.255	–0.007	–0.000
ΔP	–0.020	–0.357	0.006	0.005
ER	–0.052	–1.038	0.113	–0.020
R	0.111	–0.422	0.183	–0.004
RIP	0.016	0.456	–0.032	–0.002
ROP	–0.042	0.307	–0.011	0.009
RAX	0.566	–0.818	–0.392	–0.060

The first vector can be interpreted as a long-run real demand for money, where the income coefficient is approximately the unity, suggesting that the velocity of money circulation is stationary. In the same way, the interest rate and exchange rate seem to have a long-run effect in this relation. Thus, the first vector can be identified as:

$$M_t - P_t = RM_t = \beta_c + \beta_Y Y_t + \beta_R R_t + \beta_{ER} ER_t + \mu_1$$

By setting $\beta_Y = 1$, the later equation become: $Y_t - RM_t = \beta_c + \beta_R R_t + \beta_{ER} ER_t + \mu_1$

The second vector includes as the more relevant variables the real income, inflation and interest rate. This seems to suggest that the second vector can be interpreted as a real income relation with strong negative inflation rate effects[16]:

$$Y_t = \beta_c + \beta_{\Delta P} \Delta p + \beta_{ER} ER_t + \beta_R R_t + \mu_2$$

In order to establish the long-run relationships between the macroeconomic policy and the agricultural sector, the two other vectors have been identified as the demand for agricultural exports and the long-run adjustment of agricultural prices (input and output prices), respectively. In the second case, special attention will be paid to see if the neutrality between agricultural and non-agricultural prices holds. The neutrality hypothesis can be formulated as follows:

[16]The same relation has been identified in Juselius (1994) in the case of United Kingdom.

$$\frac{\partial\, RIP}{\partial\, ROP}\frac{RIP}{ROP} = 1$$

As the variables are in logarithms, then the cointegration parameters can be interpreted as long-run elasticities and the neutrality hypothesis is represented by:

$$\frac{\partial\, RIP}{\partial\, ROP}\frac{RIP}{ROP} = \frac{-\beta_{RIP}}{\beta_{ROP}} = 1$$

Finally, the real demand of agricultural exports is specified as a function of the exchange rate and agricultural input prices. The justification of this identification is that the exchange rate determines the competitive position of the demand of the agricultural exports vis-à-vis the movement of exchange rate. However, it should be noted that agricultural exports are sensitive to agricultural production that is also related to the input price level. This can be represented by:

$$RAX_t = \beta_c + \beta_{ER}ER_t + \beta_{RIP}RIP_t + \mu_4$$

Following Johansen and Juselius (1994), the restriction structure mentioned above can be formulated as a general hypothesis $H_\beta = (H_1\varphi_1, H_2\varphi_2, H_3\varphi_3, H_4\varphi_4)$, where the corresponding H_i matrices are given in Table 6.

Table 6. Long-run identification in the I(1) system[†]

$$H_\beta : \beta = (H_1\varphi_1, H_2\varphi_2, H_3\varphi_3, H_4\varphi_4)$$

$$\beta = \begin{pmatrix} RM \\ Y \\ \Delta P \\ ER \\ R \\ RIP \\ ROP \\ RAX \\ C \end{pmatrix} \times \begin{pmatrix} 1 & 0 & 0 & 0 \\ -1 & 1 & 0 & 0 \\ 0 & * & 0 & 0 \\ * & * & 0 & * \\ * & * & 0 & 0 \\ 0 & 0 & 1 & * \\ 0 & 0 & * & 0 \\ 0 & 0 & 0 & 1 \\ * & * & * & * \end{pmatrix}$$

Restriction matrices Hi

$$H_1 = \begin{pmatrix} 1 & 0 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix} \quad H_2 = \begin{pmatrix} 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \quad H_3 = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{pmatrix} \quad H_4 = \begin{pmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}$$

[†]An * indicates that the coefficient is not restricted.

Then, we are going to consider an alternative set of restrictions for the cointegration vectors showed in Table 6. On one hand, inflation rate is introduced as a restriction variable in the money velocity equation. Real money is introduced in the price adjustment equation as an explanatory variable of the long-run agricultural price divergence. The new sets of restrictions as well as the new Hi matrices for the identification of the cointegration space are given in Table 7.

To identify the four-cointegration vectors mentioned before, the rank condition given in (13) should be satisfied. In our case, this condition is satisfied and as the total number of restriction is 20, the system is over-identified. The likelihood ratio test statistic value was 28.78, which was above the critical values at the 5% level of significance ($\chi^2(8) = 15.5$). Then, restrictions formulate in Table 6 are rejected.

Table 7. Final long-run identification[†]

$$\beta = (H_1\varphi_1, H_2\varphi_2, H_3\varphi_3, H_4\varphi_4)$$

$$\beta = \begin{pmatrix} RM \\ Y \\ \Delta P \\ ER \\ R \\ RIP \\ ROP \\ RAX \\ C \end{pmatrix} \times \begin{pmatrix} 1 & 0 & * & 0 \\ -1 & 1 & 0 & 0 \\ * & * & 0 & 0 \\ * & * & 0 & * \\ * & * & 0 & 0 \\ 0 & 0 & * & * \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ * & * & * & * \end{pmatrix}$$

Restriction matrices Hi

$$H_1 = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ -1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \quad H_2 = \begin{pmatrix} 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \quad H_3 = \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix} \quad H_4 = \begin{pmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}$$

[†]An * indicates that the coefficient is not restricted.

This means that data not support the existence of long-run income homogeneity (money velocity cannot be stationary with the inclusion of interest rate and exchange rate). Likewise, the hypothesis of stationary price adjustments is clearly rejected. That is, both prices do not follow an equilibrium relationship and, as a consequence, the long-run neutrality adjustment of agricultural prices cannot be supported by the data. Following Tweeten (1980), these results indicate the existence of a cost-price-squeeze in the long-run and would indicate that there exist external forces that would explain the divergence between input and output prices (i.e. monetary policy).

The rank conditions are satisfied and, as we have 18 restrictions, a test for over-identification has been used. The LR statistic was 9.14, which is well under the critical value at the 5% level of significance ($\chi^2(6) = 12.6$). Then, we fail to reject the null hypothesis and the cointegration vectors are identified. The new estimated β and α matrices are shown in Table 8.

All coefficients are statistically significant and have the expected theoretical signs. The elasticity of agricultural exports with respect to the exchange rate exceeds unity (1.211). The negative and significant elasticity confirms the theoretical expectation that the exchange rate movements have important negative effects on the agricultural exports demand. Following Chamber and Just (1979), the exchange rate elasticity of exports of good (i) in equilibrium is given by $\gamma_i = \lambda_i \varepsilon_{ii}$ (where λ_i is the price elasticity with respect to exchange rate and ε_{ii} is the own-price elasticity of home country excess supply (see footnote 1). Thus, the response of exports is essentially limited by the elasticity of excess supply. If excess supply is elastic, the percentage increase in exports will exceed the percentage increase in price to exchange rate depreciation. Devadoss *et al.* (1987) obtained an elasticity of exports demand with respect to exchange rate of –1.052 while Houthakker and Magee (1969) found an elasticity of –0.96. No references exist in the Spanish case.

The other interesting equation from the agricultural point of view is the agricultural prices equation. The elasticity of producer prices with respect to input prices is 0.728 implying that given other conditions unchanged, a 1% increase in input prices would raise producer prices by 0.72%. Thus, any shock in monetary policy produces a faster adjustment in input prices than in output prices. This result suggests that farmers lost competitiveness in long-run. In other words, producers are not able to fully transmit any increase in production costs to the output prices.

Table 8. Estimated β and α matrices under long-run identification

		β matrix				Standard - errors for β			
β =	MR	1.000	0.000	0.336	0,000	0.000	0.000	0.059	0.000
	Y	-1.000	1.000	0.000	0,000	0.000	0.000	0.000	0.000
	ΔP	12.127	20.721	0.000	0,000	0.000	1.286	0.000	0.000
	ER	0.129	-0.113	0.000	1,211	0.927	0.023	0.000	0.082
	R	0.171	0.095	0.000	0,000	0.063	0.013	0.000	0.000
	IP	0.000	0.000	-0.728	0,794	0.063	0.000	0.085	0.051
	OP	0.000	0.000	1.000	0,000	0.000	0.000	0.000	0.000
	AEXP	0.000	0.000	0.000	1,000	0.000	0.000	0.000	0.000
	C	0.342	-4.067	-2.128	-7.511	0.122	0.159	0.356	0.137
		α matrix				t - values for α			
α =	MR	-0.345	0.196	0.008	0.040	-3.616	2.884	0.353	2.543
	ΔY	0.068	-0.073	0.029	0.001	2.253	-3.389	3.908	0.275
	Δ^2P	0.151	-0.142	0.064	0.023	2.330	-2.998	3.840	2.100
	ΔER	-0.987	0.781	0.255	0.063	-4.241	3.635	3.380	1.263
	ΔR	-0.735	0.581	0.123	-0.105	-1.458	1.633	0.986	-1.266
	ΔIP	0.263	-0.215	0.067	-0.074	1.989	-2.140	1.267	-2.126
	ΔOP	0.432	-0.322	0.089	0.067	1.659	-0.692	1.987	0.930
	ΔAEXP	0.542	-0.432	-0.262	-0.576	2.536	-2.025	-2.380	-7.332

Short-run dynamics

In order to identify the short-run dynamics of the variables; a structural VAR approach in error-correction form (SVECM)[17] has been considered. The SVECM can be obtained by premultiplying the reduced form of the VECM by a (p×p) A_0 matrix. The model, then, becomes:

$$A_0 \Delta Z_t = \sum_{i}^{k-1} A_i \Delta Z_{t-i} + a_\alpha(\hat{\beta}' Z_{t-1}) + \delta_1 D1 + \delta_2 D2 + u_t \qquad [11]$$

where $u_t \sim iid(0,\Omega)$; $A_i = A_0\Gamma_i$, $a_\alpha = A_0\alpha$; D1 and D2 are the dummy variables; $u_t = A_0\varepsilon_t$ and $\Omega = (A_0)\Sigma(A_0)'$.

The A_0 coefficients contain the contemporaneous linkages between all the endogenous variables in Z_t. Equation [11] can be reformulate in compact notation as:

$$AY_t + \delta_1 D1 + \delta_2 D2 = u_t \qquad [12]$$

where $A = [A_0, A_1,..., A_{k-1}, a_\alpha]$ and $Y_t = [\Delta Z_t, \Delta Z_{t-1},.., \Delta Z_{t-k}, \hat{\beta}' Z_{t-1}]'$.

The Identification of the short-run model requires the introduction of some restrictions[18] in the matrix A. In our case, we have introduced restrictions derived from the theoretical existing relationships between the macroeconomic and agricultural variables. These restrictions are of exclusion type, implying that certain variables are excluded from the relationship. The identification process requires also that the matrix A_0 have unity values in the diagonal matrix after normalization.

Johansen and Juselius (1994) formulate identifying restrictions on the columns of the A matrix in the same form as in the long-run case. The generic expression is:

$$H_A : A = (H_1\varphi_1, H_2\varphi_2,..., H_p\varphi_p) \qquad [13]$$

where H_1, H_2,..., H_p are (p×q_i) matrices; q_i is the number of unrestricted parameters in each p equations and φ_i are vectors of the corresponding estimated parameters in the system.

[17]See Hendry and Doornik (1994) for a full discussion of this approach.
[18]The model requires to be identified al most p(p-1) restrictions.

Several short-run identifications have been tread and most of them were accepted. The final identification chosen is consistent with theoretical considerations and it is not very sensitive to slight changes in the restrictions imposed. The final model has been estimated using Full Information Maximum Likelihood (FILM) procedure. The impulse response functions likelihood procedure shows the response of each variable in the system to a shock in any of the variables. For each response, the standard deviation in computed using Lültkepohl's (1993) procedure[19]. As 64 impulse response functions are obtained, in Fig. 1 only the responses of agricultural variables to a shock in the macroeconomic variables are shown which, on the other hand, is the aim of this paper.

The estimated SVECM was reparametrised to its equivalent formulation in levels and then the impulse function has been computed as in the case of a VAR(p)[20].

In general terms, results are quite consistent with what it was expected. A shock in the demand for money generates higher interest rate and then the currency is appreciated. As a consequence, agricultural exports decrease. Although not shown in the figure, the response of the exchange rate to a shock in the demand for money is almost null in the very short-run, which explains the little response of agricultural exports. Input prices do not change in the short-run while producer prices increase immediately. Higher producer prices generate a loss of competitiveness, which makes exports decrease after two quarters. What is noticeable is that contrary to the long-run, producer prices react quicker and in higher magnitude that input prices.

A shock in the exchange rate[21] (appreciation) generates an immediate reaction on the agricultural exports which almost arrives to 100% after three quarters. The effect, however, is a very short in nature as after one year it comes back to equilibrium. On the other hand, input prices decrease, as the impact on general price level is negative. In the very short-run producer prices react positively. Two main reasons would explain this reaction. Most of the annual crops have same kind of intervention. In products in which producers are able to control the short-run supply an increase in input prices reduces the short-run production and, then, prices increase. After two or three quarters, production would increase and prices go down.

Finally, the response of agricultural exports to a shock in interest rate is negative although lower in magnitude with respect to changes in the exchange rate. Input prices only increase significantly after three quarters. The response of producer prices is positive but the magnitude is lower than in the case of a shock in demand for money.

Concluding remarks

The objective of this paper was to analyse the relationships between the macroeconomic policy and the agricultural sector. Specifically, we were interested in the effects on the agricultural exports and on the relationships between producer and input prices which are the main determinants of the competitiveness of the agricultural sector. This paper is one of the first attempts to analyse such effects in Spain.

The chosen methodological approach is the specification and estimation of a VAR model. Although, most of the literature on this topic has used VAR models to analyse short-run dynamics, recent developments on time series analysis allow to distinguish between long-run and short-run effects. Eight variables have been used which capture most of the existing relationships between macroeconomics and agriculture. The existence of two I(2) variables (money supply and general price level) has complicated in a certain way the analysis. One way to proceed, as most of the existing literature, is to assume money neutrality and work with an I(1) system following the now already know Johansen procedure. However, in this paper we have formally tested for money neutrality and results indicate that this hypothesis holds.

[19]As the most interesting outcome is the impulse response function and due the space limitations, the identification procedure is not included is it is available from authors upon request.
[20]A complete description of the impulse response function computation is given in Lutkepohl (1993).
[21]Exchange rate increase means appreciation.

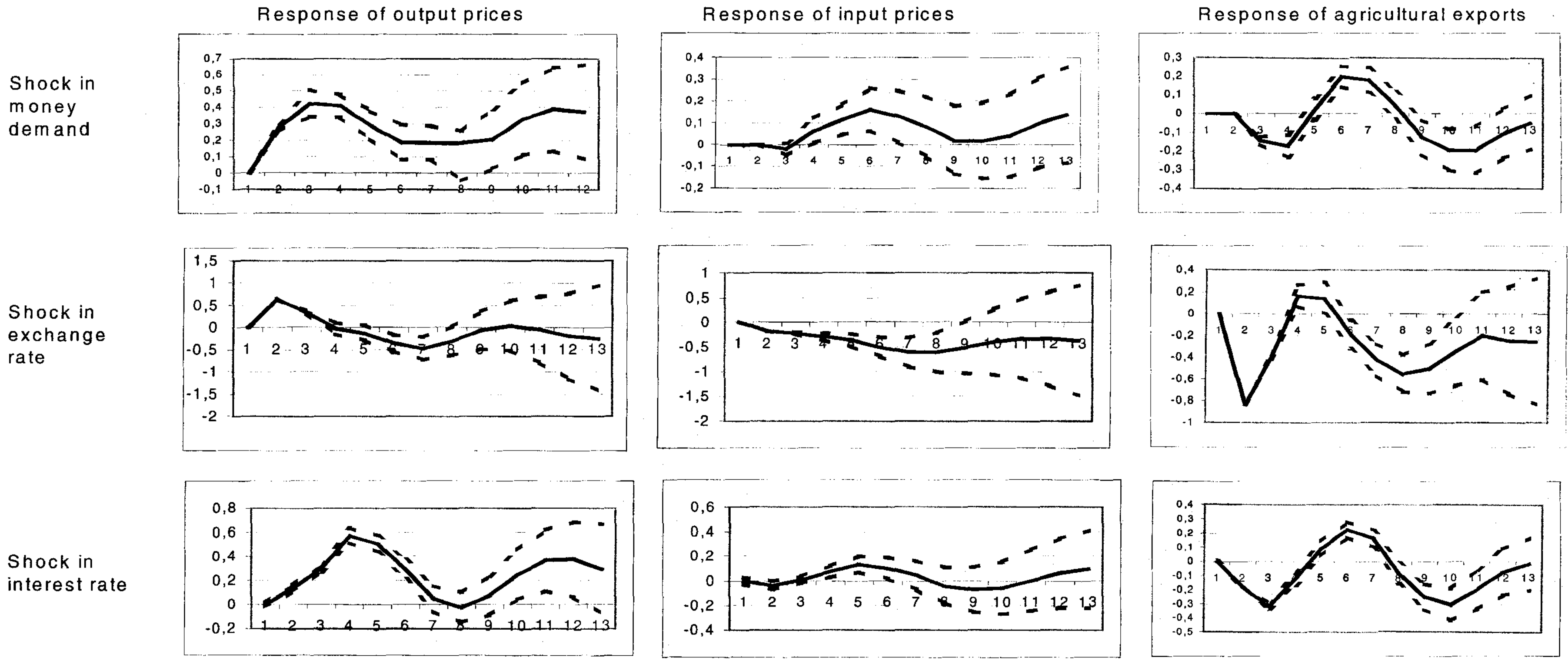

Fig. 1. Impulse response functions of agricultural variables to shock in macroeconomic variables.

A VAR system of real I(1) variables have been then specified and estimated and long-run as well as short-run analysis have been performed. In the long-run, farmers loose competitiveness as input prices react quicker and more intensively that producer prices. On the other hand, only 75% of input prices increases are translated to output prices. As far as agricultural exports concerns, the two main determinants in the long-run are exchange rate and input prices. The demand for exports is very elastic in the long-run to changes in the exchange rate. This is due to the fact that most of agricultural exports go to the European Union, which is a very competitive market as the demand is almost saturated.

In the short-run, the situation of relative agricultural prices is different. In that case, output prices are more flexible and react quickly than input prices. In this sense, this result is quite consistent with the literature as the first studies arrive to the same conclusion while in recent studies using cointegration results are similar to those obtained in this paper in the long-run. Also, in the short-run, agricultural exports are more sensitive to agricultural prices than to any macroeconomic variables. It takes around two quarters to be fully transmitted. The effects of a shock in money demand, interest rate and exchange rate are very similar on the agricultural variables.

The current trend of lower interest rate in Spain has important effects on the foreign competitiveness of the Spanish agricultural sector. On the other hand, inflation is under control so it is expected than in the future the main determinants of the competitiveness of the agricultural sector will be domestic prices and interest rate. To conclude, it has to be said that results presented in this paper depend on the variables and sample period chosen. New variables could be included in the future. In this sense, foreign direct investment could be included when data are more reliable that they are now.

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