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MACROECONOMICS AND AGRICULTURE IN TUNISIA

Monia Ben Kaabia
Departamento de Análisis Económico
Universidad de Zaragoza
Email: monia@unizar.es

José M. Gil
CREDA-UPC-IRTA
Edifici ESAB - Campus del Baix Llobregat
Av. del Canal Olímpic s/n
08860-Castelldefels (Barcelona)
Ph: ++34-935521210
Fax: ++34-935521121
e-mail: Chema.Gil@upc.edu

Houssem E. Chebbi
Sfax University (Tunisia)
e-mail: chebbihe@planet.tn



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MACROECONOMICS AND AGRICULTURE IN TUNISIA

Abstract

This paper aims to analyse the impact of changes in the monetary policy and the exchange rate on agricultural supply, prices and exports. The methodology used is based on the multivariate cointegration approach. Ten variables are considered: interest and exchange rates, money supply, inflation, agricultural output and input prices, agricultural supply and exports, income and the rate of commercial openness. Sample period covers annual data from 1967 to 2002. Due to the short-sample period, two subsystems are considered. First, long-run relationships are identified in each subsystem. Second, both subsystems are merged in order to calculate the short-run dynamics. Results indicate that changes in macroeconomic variables have an effect on the agricultural sector but the reverse effect does not hold.

Key words: Macroeconomic policy, agro-food sector, Tunisia, impulse-response functions

JEL Classification: C32, N57, O31

1. Introduction

The ongoing globalisation process in the world economy is a big challenge for Tunisia, a country which has suffered a complex process of structural economic reforms. The Adjustment Structural Program implemented in 1986 generated a new environment of economic success. All sectors of economy started to recover and exports dramatically increased being one of the main contributors to economic development. As an example, in the last five years the Tunisian GDP increased at a 5.5% annual rate while inflation was maintained around 3.5 %.

The agro-food sector in Tunisia plays an important role in Tunisian economy. It generates around 14% of total GDP, employs 22% of total labour force and agro-food exports represent around 15% of total exports, although still depending to a great extent on weather conditions. Moreover, since 1986 the agricultural sector is undergoing a modernization process characterized by a progressive intensification and the use of technology. However, the agricultural production has not been able to meet the needs of an increasing population. In general, the Government favoured imports of raw materials and food, which have provoked a progressive deterioration of trade balance. The agricultural policy was, then, oriented into two directions: 1) to promote the production of agricultural products in which self-sufficiency was low, through the implementation of a subsidies program (food security); and 2) to encourage the production of food products in which Tunisia had traditionally had a competitive advantage (olive oil, fruits, vegetables, etc) to finance the agricultural trade deficit.

In many cases, results from such policies were, to some extent, different from those expected as the effect of many macroeconomic variables (as a consequence of the Adjustment Structural Program) were not taken into account. Although not explicitly recognized, changes in the macroeconomic policy have become increasingly important for the agro-food sector, as Tunisian agriculture has become more capitalized and more dependent on international markets, thereby becoming more vulnerable to changes in interest rates, exchange rates and international growth rates.

The aim of this paper is, precisely, to provide a methodological approach taking into account data limitations to explain the relationships between macroeconomic variables and the agricultural sector in Tunisia. Special attention is paid to the distinction between long-run structural relationships and short-run dynamics. Up to our knowledge, this is the first attempt to analyse such relationships in Tunisia. The existing literature on Tunisia is quite descriptive focussing on the evolution of agricultural trade flows which are only explained by changes in the agricultural policy (Arfa, 1994; Allaya, 1995; and El Abassi, 1995, among others).

Since the mid seventies, a number of theoretical and empirical studies have analysed the impact of macroeconomic variables on the relative performance of the agricultural sector (see In and Mount, 1994, for a literature review on this topic). In the early studies, macroeconomic variables (income,

interest rate, exports,...) were introduced as purely exogenous in agricultural sector models. The paper by Schuh (1974) could be considered as the starting point of a second group of studies emphasizing the role of exchange rate in explaining agricultural variable fluctuations (Chambers and Just, 1979, 1981; Longmire and Morey, 1983; and Batten and Belongia, 1986). However, these empirical investigations neglect not only the possible effect of exchange rate changes on other macroeconomic variables (which can influence agricultural prices and exports indirectly) but also the effects of other macroeconomic variables (such as interest rates) both on exchange rate and agricultural variables. In this context, Chambers (1984) develops a general equilibrium model in order to analyse the effect of macroeconomic variables on agricultural trade where the exchange rate, income, interest rate as well as usual agricultural variables are treated as endogenous.

Finally, it is possible to identify a third group of papers dealing with the analysis of the dynamic linkages between monetary variables and the agricultural sector. The question of money neutrality in the agricultural sector, and the speed of price adjustments, has been considered of central importance for policy analysis (Bessler and Babula, 1987; Devadoss and Meyers, 1987; Taylor and Spriggs, 1989; Larue and Babula, 1994; Dorfman and Lastrapes, 1996, among others).

Results from most of the above-mentioned studies substantially differ from each other, and, in many cases, they are even contradictory. There exist alternative explanations for such differences: samples are not homogeneous, the number of variables included differs as well as their treatment as endogenous or exogenous, and the different methodological approaches used. However, there seems to exist a consensus on the fact that models analysing macroeconomic linkages to the agricultural sector should include the more relevant macroeconomic variables of the country being analysed and should treat them as endogenous (Devadoss et al., 1987; Taylor and Spriggs, 1989; Denbaly and Torgerson, 1991; Thraen et al., 1992; In and Mount, 1994; Ben Kaabia and Gil, 2000; among others). Partly for this reason, most of the analyses on this topic have recently been conducted using Vector Autoregression (VAR) models. This is also the methodological approach we have followed in this paper although adapted to take into account data limitations and their stochastic properties.

The paper is organized as follows. The data used in this study, their stochastic characteristics as well as the methodological approach are presented in Section 2. Long-run equilibrium relationships are analysed in section 3. The short-run dynamics is considered in section 4. Finally, some concluding remarks are outlined.

2. Data and methodological approach

Since the Sims' (1986) seminal paper, VAR models have been one of the most widely used analysis tools to analyse the dynamic relationships between macroeconomic and agricultural variables. In VAR models, all variables are considered endogenous and no zero/one restrictions are imposed on the variables in the system. However, recent developments in time series analysis have modified the econometric framework for analysing such relationships. The concepts of non-stationarity and cointegration have become very popular and have to be explicitly tested to properly specify an econometric model. In this new context, Johansen (1988) and Johansen and Juselius (1990, 1992 and 1994) provide an interesting methodology that allows the researcher to distinguish between the short and the long run. On the one hand, it is possible to identify the long-run structural relationships among a set of variables and how variables in the system adjust to deviations from such long-run equilibrium relationships. On the other hand, it is possible to calculate the impulse response functions in a similar way to that in the VAR models. This distinction is useful as economic restrictions are considered to be long-run in nature while it is also interesting, for the policy analysis, to know how the system adjusts to disequilibrium.

In this paper we have followed this methodological approach although we have introduced some modifications in order to adapt it to data limitations. Availability of data is a major problem for economic modelling in Tunisia. It is difficult to find a large enough sample period for many economic variables. In this study 10 variables have been considered which collect the most important information in relation to macroeconomic variables and the agricultural sector (see the Appendix for

data sources and units of measurement): 1) Real Exchange rate (ER), defined as national currency (TND) per US dollar taking into account both the US and Tunisian consumer price indices; 2) Real money supply (M) (money supply (M_2) divided by the consumer price index); 3) Interest rate (R), defined as the one-year money market interest rate; 4) Inflation (P) expressed as the Consumer Price Index in first differences; 5) Real Gross Domestic Product (GDP); 6) Real farm output prices (PP), calculated as nominal farm output prices divided by the Consumer Price Index); 7) Real farm input prices (IP)¹, calculated as nominal farm input prices divided by the Consumer Price Index); 8) Real agricultural exports (AX), calculated as the nominal exports value divided by the consumer price index); 9) Agricultural output (AP), calculated as the value of the Tunisian Agricultural Output divided by the Consumer price Index; and 10) Rate of commercial openness (RCO) calculated by dividing the international trade flows (imports + exports) by the GDP. This variable provides an indication on how the Tunisian economy is inserted in the world trade.

All variables are in logarithms, except for the interest rate and the inflation, which are in a percentage form and are divided by one hundred to make the estimated coefficients comparable with logarithmic changes. The sample period covers annual data from 1967 to 2002. Time series univariate properties have been examined by using unit root tests. As in small samples such tests have limited power (Blough, 1992), two alternative unit root tests developed by Elliot et al., (1996) and Ng and Perron (2001) as well as the stationary test from Kwiatkowski et al. (1992) (KPSS) have been applied. All tests indicated that all variables were $I(1)$ ²

Taking into account the number of variables, the number of observations available for each variable and that all variables are $I(1)$, the methodological approach followed in this paper consist of the following steps:

- i) The ten-variable system is divided into two subsystems. The first one has been defined by including: the real money supply, the inflation, the GDP, the farm input and output prices and the interest rate. Furthermore, taking into account the characteristics of the Tunisian economy, we have considered the interest rate as purely exogenous. The second subsystem includes the following seven variables: the farm input and output prices, the agricultural exports, the agricultural production, the exchange rate, the interest rate and the rate of commercial openness. Within this subsystem also the interest rate as well as the rate of commercial openness are defined as purely exogenous³.
- ii) Under the assumption of exogeneity for certain variables, the multivariate cointegration procedure developed by Pesaran et al. (2000) is used to test for cointegration in both subsystems. Moreover, cointegration vectors are identified as long-run meaningful economic relationships.
- iii) Merging results from the two subsystems into a single system with the original 10 variables, impulse response functions are computed to analyse short-run dynamics and to test the exogeneity assumptions made in the first step.

3. Long-run analysis

3.1. Model specification and cointegration rank

All variables in each subsystem were $I(1)$ and, then, a Vector Error Correction Model has been specified for each subsystem. The methodology developed by Pesaran et al. (2000) is used to determine the cointegration rank. These authors modified the Johansen (1988) procedure to explicitly allow for the introduction of exogenous variables. The base-line econometric specification for multivariate cointegration is a VAR(p) representation of a k-dimensional time series vector Y_t reparametrized as a Vector Error Correction Model (VECM):

¹ Fertilizer prices are used as a proxy in this study.

⁵ Results are not shown due to space limitations. They are available upon request.

³ In a further sep in the modelling process, specific tests will be carried out to test for the exogeneity of the mentioned variables.

$$\Delta Y_t = \mu D_t + \Gamma_1 \Delta Y_{t-1} + \dots + \Gamma_{p-1} \Delta Y_{t-p+1} - \Pi Y_{t-1} + e_t \quad (1)$$

where, Y_t is a $(k \times 1)$ column vector of variables; D_t is a vector of deterministic variables (intercepts, trend...) where μ is the matrix of parameters associated with D_t ; Γ_i are $(k \times k)$ matrices of short-run parameters ($i=1, \dots, p-1$), where p is the number of lags; Π is a $(k \times k)$ matrix of long-run parameters and e_t is the vector of disturbances $\text{iid}(0, \Sigma)$.

When exogenous variables are considered the vector Y_t vector can be partitioned as $Y_t = (Z_t', X_t')$, where Z_t is a $(m \times 1)$ vector of endogenous variables and X_t is a $(n \times 1)$ vector of exogenous variables ($n=k-m$), which can be considered as the "long-run forcing" variables in the system, that is, changes in X_t have a direct influence on the variables Z_t , while they are not affected either by the changes in the equilibrium relationships nor by past changes in Z_t . This is equivalent to the notion that the set of variables Z_t do not Granger-cause X_t . According to the mentioned partition the error term e_t can be decomposed as follows:

$$e_t = (e'_{yt}, e'_{xt})'$$

with covariance matrix given by:

$$\Omega = \begin{pmatrix} \Omega_{yy} & \Omega_{yx} \\ \Omega_{xy} & \Omega_{xx} \end{pmatrix} \quad (2)$$

According to (3), the error terms of the endogenous variables (e_{yt}) can be represented in terms of the e_{xt} as follows:

$$e_{yt} = \Omega_{yx} \Omega_{xx}^{-1} e_{xt} + u_t \quad (3)$$

where, $u_t \sim \text{IN}(0, \Omega_{uu})$ and $\Omega_{uu} = \Omega_{yy} - \Omega_{yx} \Omega_{xx}^{-1} \Omega_{xy}$, being u_t independent of e_{xt} .

Substituting (3) in (1) and considering a similar partition for the other matrices, $\mu = (\mu'_z, \mu'_x)'$, $\Pi = (\Pi'_z, \Pi'_x)'$, $\Gamma_i = (\Gamma'_{zi}, \Gamma'_{xi})'$, ($i=1, 2, \dots, p-1$), we get a conditional model for ΔZ_t as a function of Y_{t-1} , ΔX_t , ΔY_{t-1} , $\Delta Y_{t-2}, \dots$, which adopts the following expression::

$$\Delta Z_t = \delta D_t + \Lambda \Delta X_t + \sum_{i=1}^{p-1} \Psi_i \Delta Y_{t-i} + \Pi_{yy,x} Y_{t-1} + u_t \quad t=1, 2, \dots, T \quad (4)$$

being : $\delta = \mu_z - \Omega_{zx} \Omega_{xx}^{-1} \mu_x$

$$\Lambda = \Omega_{zx} \Omega_{xx}^{-1}$$

$$\Psi_i = \Gamma_{zi} - \Omega_{zx} \Omega_{xx}^{-1} \Gamma_{xi} \quad i=1, 2, \dots, p-1$$

$$\Pi_{zz,x} = \Pi_z - \Omega_{zx} \Omega_{xx}^{-1} \Pi_x$$

If variables in X_t are not cointegrated, that is, $\Pi_x = 0$ (and, then, $\Pi_{zz,x} = \Pi_z$), Pesaran et al. (2000) show that the k -variable system defined in (1) can be decomposed in the following two subsystems:⁴

- Conditional subsystem: $\Delta Z_t = \delta D_t + \Lambda \Delta X_t + \sum_{i=1}^{p-1} \Psi_i \Delta Y_{t-i} + \Pi_z Y_{t-1} + u_t \quad (5)$

⁴ Under such decomposition, variables in X_t are assumed to be weakly exogenous with respect to the cointegration space. Moreover, if variables in Z_t do not Grange-cause X_t , then such variables are assumed to be strongly exogenous with respect to such cointegration space, that is, they would be only explained by their own past in the marginal subsystem.

- Marginal subsystem:
$$\Delta X_t = \mu_x + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta Y_{t-i} + \varepsilon_{xt} \quad (6)$$

Taking into account equations (5) and (6), to test for cointegration is equivalent to test for the Rank (r) of the matrix Π_z :

$$H_r : \text{Rank} [\Pi_z] = r \quad r=0, \dots, m \quad (7)$$

To test for the number of cointegrating vectors (r), Pesaran et al., (2000), following Johansen (1988), proposed two statistics: the trace statistic and the λ_{\max} statistic. If the hypothesis of cointegration is not rejected ($0 < r < m$), Y_t is said to be cointegrated in the sense that there exists a $k \times r$ matrix β such that $(\beta' Y_{t-1})$ is stationary and, consequently, the cointegration relationships can be formally expressed as $\Pi_z = \alpha_z \beta'$. This procedure has been applied to the two subsystems described in the last section. Both subsystems are estimated including two lags⁵ and a constant restricted to the cointegration space⁶. Multivariate tests for autocorrelation (Godfrey, 1988) and normality (Doornik and Hansen, 1994) have been carried out to check for model statistical adequacy before applying the reduced rank tests. Results indicated that both subsystems could be considered correctly specified⁷.

Table 1 shows the results from cointegration tests in both subsystems. As can be observed, for the first subsystem (upper part of Table 1) results from the λ -max and the trace tests indicate that there exist two cointegration vectors among the six variables includes while for the second subsystem (lower part of Table 1) results differ depending on the level of significance (two and three cointegration vectors for 5 and 10% levels of significance, respectively).

Table 1. Results from cointegration tests

First subsystem $Y' = \{M, P, GDP, PP, IP, R\}$ ¹							
$H_0: r$	$H_a: p-r$	λ -max.	Trace	Critical values λ -max ²		Critical values Trace ²	
				(10%)	(5%)	(10%)	(5%)
0	5	46,88	124,76	34,99	37,48	82,17	86,58
1	4	37,72	77,88	29,01	31,48	59,07	62,75
2	3	20,92	38,15	22,98	25,54	39,12	42,40
3	2	11,37	19,23	16,74	18,88	22,76	25,23
4	1	7,85	7,85	10,50	12,45	10,50	12,45
Second subsystem $Y' = \{PP, IP, AX, AP, ER, R, RCO\}$ ¹							
$H_0: r$	$H_a: p-r$	λ -max.	Trace	Critical values λ -max ²		Critical values Trace ²	
				(10%)	(5%)	(10%)	(5%)
0	5	42,84	123,36	37,81	40,57	92,93	97,57
1	4	34,91	80,15	32,00	34,69	67,83	72,15
2	3	28,42	47,06	26,08	28,49	45,89	49,43
3	2	16,17	23,85	19,67	21,92	27,58	30,46
4	1	7,68	7,68	13,21	15,27	13,21	15,27

¹ See the Appendix for variable definitions

² Critical values are taken from Pesaran et al. (2000).

Taking into account the relatively large dimension of the VECM and the small sample available, the outcome of the test procedure has to be interpreted with some caution. Several simulation studies show (Abadir et al., 1999; Gredenhoff and Jacobson, 2001; and Johansen and Juselius, 2000) that the asymptotic critical values may not be very close approximations in small samples. For that reason, we

⁵ A small-sample adjusted Likelihood Ratio statistic has been used considering a maximum lag of three periods taking into account the sample size.

⁶ Results from unit root tests indicated that almost all the variables were non-stationary with non zero-means.

⁷ Results from multivariate first-order autocorrelation tests were 21.14 and 23.85 for the first and the second subsystem, respectively, which were well below the critical value at the 5% level of significance ($\chi_{25}^2 = 37.65$). Results from multivariate normality tests were 14.43 and 17.83, for the first and the second subsystem, respectively, which were well below the critical value at the 5% level of significance ($\chi_{10}^2 = 18.31$).

have also studied the roots of the companion matrix and the t-ratios of the α_z parameters from the last cointegration vector (Juselius, 1995). For both subsystems, all the roots were inside the unit circle, indicating that all variables were I(1). Moreover, the eigenvalues of the companion matrix show that, for both subsystems, the first four roots were close to unity while the rest were quite small. In other words, we could not reject the null of two and three cointegrating vectors for the first and the second subsystems, respectively. Finally, all t-ratios of the α_z parameters of the third cointegration vector for the first subsystem were not significant, while in the second subsystem, some of them were significant⁸. Thus, the first subsystem has been specified with two cointegrating vectors, whereas three cointegration vectors have been chosen for the second one.⁹

3.2 Long-run structural relationships

Identifying economically interpretable relations is the primary aim of this analysis. However, Juselius (1994) argues "the interpretation of the unrestricted cointegration space is far from straightforward when there are more than one cointegrating vector". Moreover, Johansen and Juselius (1994) suggest that only sometimes the unrestricted cointegrating vectors, surprisingly, can be directly interpreted in terms of theoretical economic relationships. Thus, some restrictions are needed in order to obtain a structural representation of such relationships.

First subsystem

Taking into account the variables included in the model as well as the economic theory which relates those variables, the following hypothetical cointegration relations could be expected:

i) A money demand equation in real terms in which the monetary aggregate is related to the inflation in Tunisia, the Gross Domestic Product and an opportunity cost represented by the interest rate:

$$(\beta_1^{\text{sys1}})' Y_t : RM_t = \beta_{\text{GDP}}^1 \text{GDP}_t + \beta_R^1 R_t + \beta_P^1 P_t + \mu^1 + \varepsilon_{1t} \quad (8)$$

It is expected that $\beta_{\text{GDP}} > 0$; $\beta_R < 0$ and $\beta_P < 0$. If $\beta_{\text{GDP}} = 1$ Equation (8) would be consistent with the Quantity Theory of Money, whereas $\beta_P = 0$ would exclude inflation to play a role in the demand for money in Tunisia.

ii) A price transmission equation:

$$(\beta_2^{\text{sys1}})' Y_t : IP_t = \beta_{\text{PP}}^2 PP_t + \mu^2 + \varepsilon_{2t} \quad (9)$$

from which it is possible to test the homogeneity condition: $-\frac{\beta_{\text{IP}}}{\beta_{\text{PP}}} = \frac{\partial \text{IP}}{\partial \text{PRC}} \frac{\text{PP}}{\text{IP}} = 1$

The two equations can be written more compactly as: $\beta'_{\text{sys1}} Y_{t-1} = \varepsilon_t \sim I(0)$

$$\text{where: } \beta^{\text{sys1}} = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} \quad (10)$$

In this paper, a two-step procedure is going to be used in order to check if (10) is supported by data. In the first step, each single restricted relation (8)-(9) is tested for stationarity leaving the other relations unrestricted. In other words, if restrictions imposed are compatible with a stationary relationship. The second step involves jointly considering the full identification of the two relationships. Juselius (1998) points out that this approach maximizes the chance of finding a correct full identification of long-run relations.

Hypotheses related to the first step adopt the general form $H_{0i}: \beta = (H_i \phi, \omega)^{10}$. In such an expression, restrictions to be tested are only placed in a single cointegration vector while the

⁸ Results are not presented due to space limitations but they are available from authors upon request.

⁹ The unrestricted cointegration space is not presented due to space limitations. Test carried out on the long-run parameters in β indicated that all of them were significant.

¹² See Johansen and Juselius (1992) for a full description of the procedure to formulate and test such hypotheses.

remaining (r-1) vectors are considered unrestricted. Johansen and Juselius (1992) suggest that this test can be used when we wish to test if there exists some vector in the cointegration space that linearly combines the variables in a particular hypothesized stationary relationship. Several hypotheses have been considered and tested. The specification of such hypotheses, as well as main results found are shown in Table 2. With respect to the first relationship, three different hypotheses have been tested. In the first one (H_{01}^{sys1}), it is tested that real money is cointegrated with interest rate, GDP and inflation, imposing also income homogeneity. Results from the Likelihood Ratio (LR) statistic indicate that the null cannot be rejected¹¹. In the second hypothesis (H_{02}^{sys1}), an additional restriction is considered ($\beta_p^1=0$). This hypothesis is strongly rejected, which means that the monetary authority is not fixing the monetary policy taking into account an aggregate money stock. The third hypothesis (H_{03}^{sys1}) is similar to the first one but excluding the income homogeneity. Also in this case, we fail to reject the null hypothesis. Finally, in relation to the second relationship, hypothesis (H_{04}^{sys1}) tests for price homogeneity in the agricultural sector. The LR statistic is under the critical value suggesting that monetary policy has a neutral effect on the real food-based prices. This means that, in the long run, input prices and output prices react in the same way and magnitude to changes in money supply.

Once it has been checked that each single equation is a cointegrated relationship, the second step consists of testing a full identification of the structural long-run relationships following Johansen and Juselius (1994). Taking into account results showed in the upper part of the Table 2, two hypotheses have been tested. The first one (H_{05}^{sys1}) jointly tests the hypotheses H_{01}^{sys1} and H_{04}^{sys1} , whereas the second tests the hypotheses H_{03}^{sys1} and H_{04}^{sys1} . Only in the second case we fail to reject the null hypothesis (the LR statistic is 11.55, which is well under the critical value at the 1% level of significance ($\chi^2(5) = 15.09$)), indicating that in Tunisia, the inflation plays a significant role in the demand for money and that agricultural prices satisfy the homogeneity condition.

Second subsystem

In the second subsystem, taking into account the variables included and results obtained in the first one in relation to the agricultural prices, the following hypothetical cointegration relations could be expected:

i) As the agricultural prices are also included in the second subsystem, and in order to check for data consistency, the first cointegration relationship would attempt to relate agricultural prices under the homogeneity restriction:

$$(\beta_1^{sys2})' Y_t : IP_t = \beta_{PP}^1 PP_t + \mu^1 + \varepsilon_{1t} \quad (11)$$

ii) The second relationship is going to be associated with an agricultural export equation for Tunisia, which would depend on the exchange rate, farm output prices and the rate of commercial openness:

$$(\beta_2^{sys2})' Y_t : AX_t = \beta_{PP}^2 PP_t + \beta_{ER}^2 ER_t + \beta_{RCO}^2 RCO_t + \mu^2 + \varepsilon_{2t} \quad (12)$$

iii) The last relationship is defined as an agricultural supply equation in which farm input and output prices, the interest rate and the rate of commercial openness are included as main potential determinants:

$$(\beta_3^{sys2})' Y_t : AP_t = \beta_{IP}^3 IP_t + \beta_R^3 R_t + \beta_{RCO}^3 RCO_t + \mu^3 + \varepsilon_{3t} \quad (13)$$

Equations (11), (12) and (13) can be formulated in compact form as:

¹¹ Several authors such as Reimers (1992) and Abadir et al. (1999) pointed out the tendency of likelihood ratio tests to over-reject in small samples when testing for the cointegration rank. Garratt et al. (1999) undertook a bootstrapping exercise to obtain critical values for testing the over-identification restrictions. The resulting critical values were higher than the asymptotic ones. This result would imply that the over-identification restrictions tested here are not rejected with higher p-values.

$$\beta^{sys2} Y_{t-1} = \varepsilon_t \sim I(0) \text{ where } \beta^{sys2} = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ 0 & * & 0 & 1 & 0 & * & * & * \end{bmatrix} \quad (14)$$

Table 2. Hypothesis restrictions tests on the cointegration vectors in the first subsystem ^a

Hypotheses on a single cointegration vector		
Hypothesis formulation	Statistic	Critical Value (1%)
$H_{01} : \beta_1^{sys1} = (\beta_1, \Phi) = (H_1\phi, \Phi)$		
$H_{01}^{sis1} : \beta_1^{sys1} Y_t = \begin{bmatrix} 1 & * & -1 & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 8,77$	9,21
$H_{02}^{sis1} : \beta_2^{sys1} Y_t = \begin{bmatrix} 1 & 0 & -1 & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(3) = 12,37$	11,34
$H_{03}^{sis1} : \beta_3^{sys1} Y_t = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 5,58$	6,63
$H_{04}^{sis1} : \beta_4^{sys1} Y_t = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(4) = 10,12$	13,28
Hypotheses on the full system		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{01} : \beta_1^{sys1} = (\beta_1, \beta_2) = (H_1\phi_1, H_2\phi_2)$		
$H_{05}^{sis1} : \beta_5^{sys1} Y_t = \begin{bmatrix} 1 & * & -1 & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} Y_t$	$\chi^2(6) = 35,16$	16,81
$H_1' = \begin{bmatrix} 1 & 0 & -1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$ and $H_2' = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		
$H_{06}^{sis1} : \beta_6^{sys1} Y_t = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} Y_t$	$\chi^2(5) = 11,55$	15,09
$H_1' = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$ and $H_2' = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		

^a $Y' = \{M, P, GDP, PP, IP, R\}$. An * indicates that the coefficient is unrestricted.

In order to test restrictions on the cointegration space, a similar approach to that mentioned for the first subsystem has been followed. However, in this case, as there are three cointegration vectors, one further step has been included. As a first step, we have carried out some tests on each individual long-run relationship leaving the rest unrestricted. The first hypothesis (H_{01}^{sys2}), as mentioned above, only tries to guarantee consistency of data used. Thus, we have tested if agricultural prices homogeneity is stationary. Results from the LR test indicate that the null cannot be rejected, the same results as in the first subsystem. Three alternative hypotheses have been defined for the agricultural exports equation. The first one (H_{02}^{sys2}), tests for a stationary relationship among agricultural exports, farm output prices and the exchange rate and the rate of commercial openness. The second one (H_{03}^{sys2}) excludes the rate of commercial openness and includes the agricultural supply. Finally, the third one (H_{04}^{sys2}) excludes the rate of commercial openness without including any other variable. Results from the LR tests indicate that only the two first hypotheses are supported by the data. Finally, in relation to the agricultural supply equation, two alternative hypotheses have been considered. In the first one (H_{05}^{sys2}) agricultural output is defined as a function of farm input prices, the interest rate and the rate of

commercial openness. In the second one (H_{06}^{sys2}) the last variable is excluded from the equation. In this case, only the first hypothesis is supported by the data.

As a second step in the identification process of the long-run relationships, we have tested restrictions on two cointegration vectors taking into account the results obtained above. Thus, three further hypotheses have been tested. In all of them, we have maintained the price homogeneity restriction. The first one (H_{07}^{sys2}) jointly tests H_{01}^{sys2} and H_{02}^{sys2} , whereas the second one (H_{08}^{sys2}) tests H_{01}^{sys2} and H_{03}^{sys2} and, finally, the hypothesis H_{09}^{sys2} jointly tests H_{01}^{sys2} and H_{05}^{sys2} . Results are shown in the middle part of Table 3. As can be observed, we fail to reject hypotheses H_{07}^{sys2} and H_{09}^{sys2} , indicating that we have a potential identification for the three-equation cointegration space. At the bottom of Table 3, results from jointly testing the hypotheses H_{01}^{sys2} , H_{03}^{sys2} and H_{05}^{sys2} are shown. The null cannot be rejected, indicating that the cointegrating space is identified.

Finally, Table 4 shows the estimated parameters of the β and α_z matrices corresponding to the two subsystems. In the first case, the two cointegrating vectors have been normalised by the real money supply and the farm out prices. In the second subsystem, the three cointegrating vectors have been normalised by the farm output prices, the agricultural exports and the agricultural supply, respectively. All coefficients are statistically significant and have the expected theoretical signs. The only exception is the positive sign of inflation in the money demand equation. This result, following Sriram (1999) would indicate that in Tunisia an expectation of increasing inflation would determine economic agents to accumulate money stock to increase nominal income. It is also interesting the sign associated to the rate of commercial openness in the agricultural export and production equations. In the first case the sign is positive as expected. In the second equation the sign is negative suggesting that a higher rate of commercial openness would generate increasing imports of food products in which Tunisia has not competitive advantages (cereals, beef, vegetable oils, etc.), negatively affecting domestic production.

The magnitude of coefficients cannot be interpreted as Juselius (1999) points out that "it is no longer possible to interpret a coefficient in a cointegrating relation as in conventional regression context....In multivariate cointegration analysis all variables are stochastic and a shock to one variable is transmitted to all variables via dynamics of the system until the system has found its new equilibrium position". On the other hand, in this type of analysis it is also convenient to consider the estimated $\alpha_{i,j}^z$ (i indicates the row and j the column) parameters as they provide valuable information about the speed of adjustment of each variable towards the long-run equilibrium. As the relationships between macroeconomic variables and the agricultural sector are of interest for this study, let us focus on such relationships.

In relation to the first subsystem, and only considering the money demand and the price equations, the first conclusion is that there seems to exist a feedback relationship between macroeconomic variables and the agricultural sector. In fact, any shock in the money demand generates a response of input and output prices. On the other hand, any change in the long-run relationship between agricultural prices affects both the income (Gross Domestic Product) and the inflation. In relation to the price transmission mechanism, although in the long-run homogeneity holds, in the short-run the situation looks different. The α_z parameters corresponding to the first cointegrating relationship indicate that input prices react quicker than output prices. This result suggests a cost-push transmission mechanism within the Tunisian agricultural sector, which is also confirmed when observing results from the second subsystem as $\alpha_{21} > \alpha_{11}$, $\alpha_{22} > \alpha_{12}$ and $\alpha_{23} > \alpha_{13}$.

In the short run there seems not to exist a close relationship between agricultural supply and exports (the α_{42} and α_{33} parameters are not significant). This result would indicate that, in Tunisia, agricultural exports depend more on other factors rather than on the agricultural production as, for example, the commercial agreements (most of the exported food products are addressed to the European Union and are subject to contingents) or decisions made by existing exporters lobbies in the most important exporting goods (olive oil, dates, citrus, etc.). In other words, agricultural policy is more oriented to support agricultural prices and producers and consumers income than to incentive

trade competitiveness. Moreover the parameter α_{31} is not significant, indicating that there is not any significant relationship between farm output prices and agricultural exports, which reinforces the idea of dissociation between agricultural exports and supply.

Table 3. Hypothesis restrictions tests on the cointegration vectors in the second subsystem^a

Hypotheses on a single cointegration vector		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{01}^{sis2} : \beta_1^{sys2} = (\beta_1, \Phi) = (H\phi, \Phi)$		
$H_{01}^{sis2} : \beta_1^{sys2} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(4) = 7,87$	13,28
$H_{02}^{sis2} : \beta_2^{sys2} Y_t = \begin{bmatrix} * & 0 & 1 & 0 & * & 0 & * & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 0,01$	6,63
$H_{03}^{sis2} : \beta_3^{sys2} Y_t = \begin{bmatrix} * & 0 & 1 & * & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 0,50$	6,63
$H_{04}^{sis2} : \beta_4^{sys2} Y_t = \begin{bmatrix} * & 0 & 1 & 0 & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 9,83$	9,21
$H_{05}^{sis2} : \beta_5^{sys2} Y_t = \begin{bmatrix} 0 & * & 0 & 1 & 0 & * & * & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 6,06$	6,63
$H_{06}^{sis2} : \beta_6^{sys2} Y_t = \begin{bmatrix} 0 & * & 0 & 1 & 0 & * & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 10,70$	9,21
Hypotheses on two cointegration vectors		
Hypothesis formulation:	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{sys2} = (\beta_1, \beta_2, \Phi) = (H_1\phi_1, H_2\phi_2, \Phi)$		
$H_{07}^{sis2} : \beta_7^{sys2} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 8,38$	15,09
$H_{08}^{sis2} : \beta_8^{sys2} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & * & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 17,23$	15,09
$H_{09}^{sis2} : \beta_9^{sys2} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ 0 & * & 0 & 1 & 0 & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 10,07$	15,09
Hypotheses on the full system		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{sys2} = (\beta_1, \beta_2, \beta_3) = (H_1\phi_1, H_2\phi_2, H_3\phi_3)$		
$H_{010}^{sis2} : \beta_{10}^{sys2} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ 0 & * & 0 & 1 & 0 & * & * & * \end{bmatrix} Y_t$	$\chi^2(6) = 15,23$	16,81
$H'_1 = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}; H'_2 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}; H'_3 = \begin{bmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		

^a $Y^* = \{PP, IP, AX, AP, ER, R, RCO\}$. An * indicates that the coefficient is unrestricted.

However, simple considering the magnitude of adjustments to long-run relationships is not enough. It is also important to look at the time path of the reactions. The impulse response functions provide relevant evidence. They are analysed in the next section.

Table 4. Estimated β and α_z matrices under long-run identification for both subsystems

First subsystem	
$\beta' = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,014 & 0,586 \\ & (0,413) & (0,026) & & & (0,007) & (0,231) \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,207 \\ & & & & & & (0,471) \end{bmatrix} \times$	$\alpha_z = \begin{bmatrix} M & -0,123 & -0,014 \\ P & (-1,827) & (-0,522) \\ GDP & 0,156 & 0,034 \\ & (4,343) & (3,752) \\ PP & 0,204 & 0,066 \\ & (3,034) & (3,883) \\ IP & 0,096 & -0,023 \\ & (1,945) & (-2,787) \\ R & 0,314 & 0,076 \\ & (1,978) & (2,691) \\ Const. & & \end{bmatrix}$
Second subsystem	
$\beta' = \begin{bmatrix} 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,269 \\ & & & & & & & (0,058) \\ 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ & (0,384) & & & (0,083) & & (0,194) & (0,823) \\ 0,000 & 0,871 & 0,000 & 1,000 & 0,000 & 0,074 & 0,894 & -0,277 \\ & (0,089) & & & (0,080) & (0,178) & (0,862) & \end{bmatrix} \times$	$\alpha_z = \begin{bmatrix} PP & -0,146 & 0,079 & -0,049 \\ IP & (-2,279) & (2,781) & (-2,083) \\ AX & 0,260 & 0,109 & -0,361 \\ & (4,014) & (4,328) & (-3,871) \\ AP & -0,217 & -0,363 & 0,400 \\ & (-0,832) & (-3,898) & (0,786) \\ ER & 0,145 & 0,260 & -0,128 \\ & (2,356) & (1,323) & (-1,932) \\ R & -0,048 & -0,028 & 0,003 \\ & (-0,723) & (-1,198) & (0,873) \\ RCO & & & \\ Const. & & & \end{bmatrix}$

Note: Values in parentheses correspond to standard deviations, in the case of the β parameters, and to t-ratios, in the case of the α_z parameters

4. Short-run dynamics

Once the VECM has been estimated, short-run dynamics can be examined by considering the impulse response functions (IRF). These functions show the response of each variable in the system to a shock in any of the other variables. The IRF are calculated from the Moving Average Representation of the VECM (see Lütkepohl, 1993 and Pesaran and Shin, 1998):

$$Y_t = \sum_{i=0}^{\infty} B_i \varepsilon_t$$

where matrices B_i ($i=2, \dots, n$) are recursively calculated using the following expression:

$B_n = \Phi_1 B_{n-1} + \Phi_2 B_{n-2} + \dots + \Phi_k B_{n-p}$; $B_0 = I_p$; $B_n = 0$ for $n < 0$; $\Phi_1 = I + \Pi + \Gamma_1$; and $\Phi_i = \Gamma_i - \Gamma_{i-1}$ ($i=2, \dots, p$). Following Pesaran and Shin (1998) the scaled Generalized Impulse Response Functions (GIRF) of variable Y_i with respect to a standard error shock in the j^{th} equation can be defined as:

$$\text{GIRF}(Y_{it}, Y_{jt}, h) = \frac{e_i' B_h \Sigma e_j}{\sqrt{\sigma_{jj}}}; \quad h = 0, \dots, n$$

where e_{is} ($s=i, j$) is the s^{th} column of the identity matrix.

The GIRF are unique and do not require the prior orthogonalisation of the shocks (reordering of the variables in the system). On the other hand, the GIRF and the orthogonalised IRF (Cholesky) coincide if the covariance matrix, Σ , is diagonal and $j=1$. Standard deviations of impulse responses are obtained following Pesaran and Shin (1998).

To analyse the short-run dynamics, as mentioned in section 2, we have first integrated the 10 variables in a full system including all restrictions on the long-run parameters shown in the previous table. As the price transmission long-run relationship was specified in the two subsystems, only four cointegration relationships were defined. Moreover, when defining the full system two additional tests were carried out. The first one was to check if the variables we considered as purely exogenous (R and RCO) they actually were. In this case, we have assumed that both variables followed a univariate

autoregressive process (i.e. they are not influenced by past values of the other variables in the system) and we have tested for the significance of the appropriate Γ_1 parameters corresponding to such equations. The value of the statistic was 31,02, which was under the critical value ($\chi^2_{17} = 33.41$) at the 1% level of significance). Second, all adjustment coefficients (α_z parameters), which were non-significant in Tables 3 and 4, were restricted to zero. The test indicated that the null was not possible to be rejected (the LR statistic was 58.92 which was under the critical value of a $\chi^2(43)=66.95$, at the 1% significance level). The final estimated model is given by:

$$\beta'Y_t = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,014 & 0,000 & 0,586 \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,207 \\ 0,000 & 0,000 & 0,000 & 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ 0,000 & 0,000 & 0,000 & 0,000 & -0,871 & 0,000 & 1,000 & 0,000 & -0,074 & 0,894 & -0,277 \end{bmatrix} \begin{matrix} M \\ P \\ GDP \\ PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \\ Const. \end{matrix}$$

$$\alpha = \begin{bmatrix} 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0333 & 0,0141 & 0,0000 & 0,0000 \\ 0,0516 & 0,0252 & 0,0000 & 0,0000 \\ 0,0521 & 0,0324 & 0,0128 & 0,0117 \\ 0,1786 & 0,1035 & 0,0464 & 0,0409 \\ 0,0000 & 0,0000 & 0,1454 & 0,0000 \\ 0,0000 & 0,1555 & 0,0000 & 0,0495 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \end{bmatrix} \begin{matrix} M \\ P \\ GDP \\ PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \end{matrix}$$

As 100 impulse-response functions are obtained, we will analyse only the estimated impulse responses of agricultural variables to a shock in the main variables in the system. Significant responses are marked with a circle. In general terms, most of the responses are not significant although they show the expected signs. In any case, as we have annual data, we cannot expect responses longer than one or two years. Figure 1 shows the responses to a shock in the real quantity of money, (through an unexpected increase in the nominal quantity of money). As expected, an expansive monetary policy affects positively to inflation, although the effect is only significant during the second year after the shock. Moreover, the increasing access to credits stimulates economic growth (the GDP increase) as well as agricultural exports. The effect on farm output prices is positive but non significant. The aim of the Tunisian agricultural policy is to support farmers' income through intervention prices, but compatible with inflation control. Agricultural prices are not allowed to increase over expected inflation in order to guarantee consumers' access to basic foods. Limited increases of agricultural prices do not stimulate agricultural production neither the demand for inputs. As a result, the impact on input prices is not significant. Finally, following the Keynesian theory, an expansive monetary policy induces exchange rate depreciation, which, on the other hand, leads to an increase of agricultural exports.

Figure 2 shows the responses to a shock in farm output prices to the most relevant variables within the system. In general terms, agricultural variables do not have any effect on macroeconomic variables and, then, such effects are not included in the Figure. Two main results are found. First, a positive increase of producer prices generates a positive response of agricultural production. The response is significant for two years. Second, it also generates an immediate response of input prices. The magnitude of such a response is higher than in the case of output prices, which is consistent with comments in section 3 about the α_z parameters. The effect on inflation is positive but not significant. As mentioned above, the government has traditionally controlled agricultural price increases to be compatible with the inflation rate in Tunisia. Finally, the effect on agricultural exports is not

significant either. Prices are not the main source of competitiveness for Tunisian agricultural exports as they are mainly subject to contingents. Moreover, the traditional policy used by the Tunisian Government to promote exports has been via exchange rate.

Figure 1. Responses to a shock in the money supply

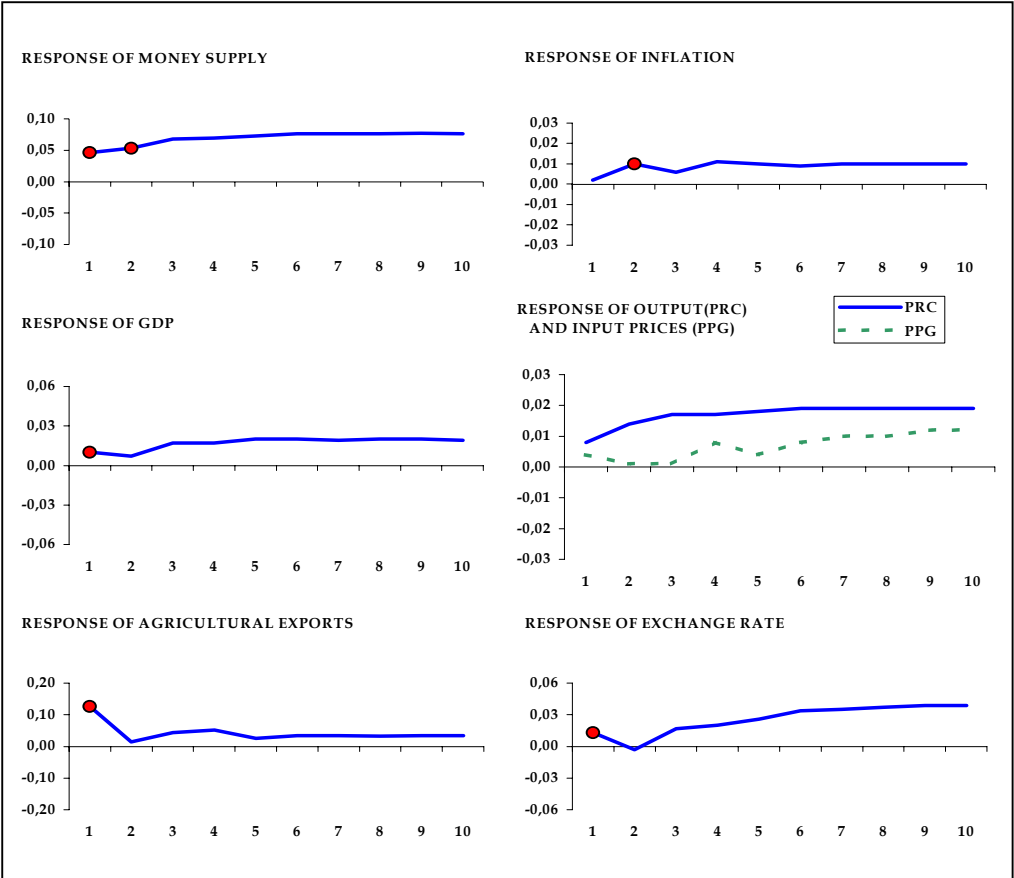
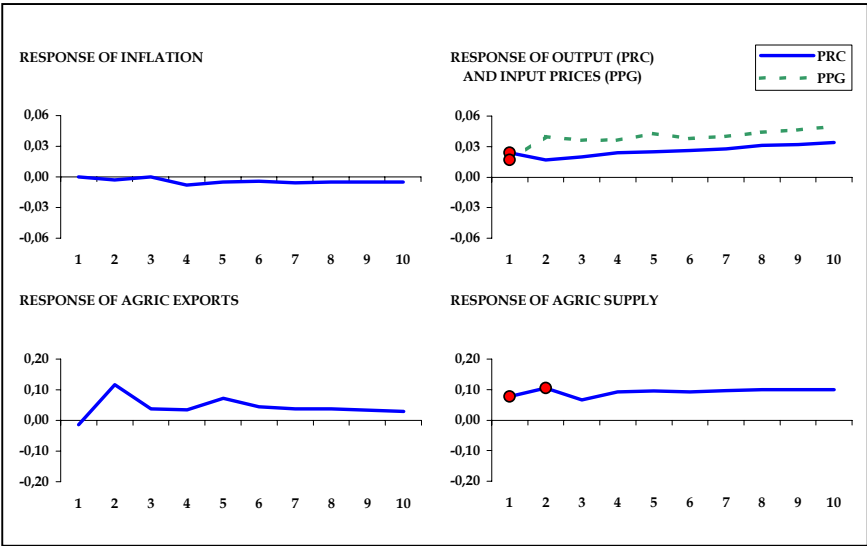


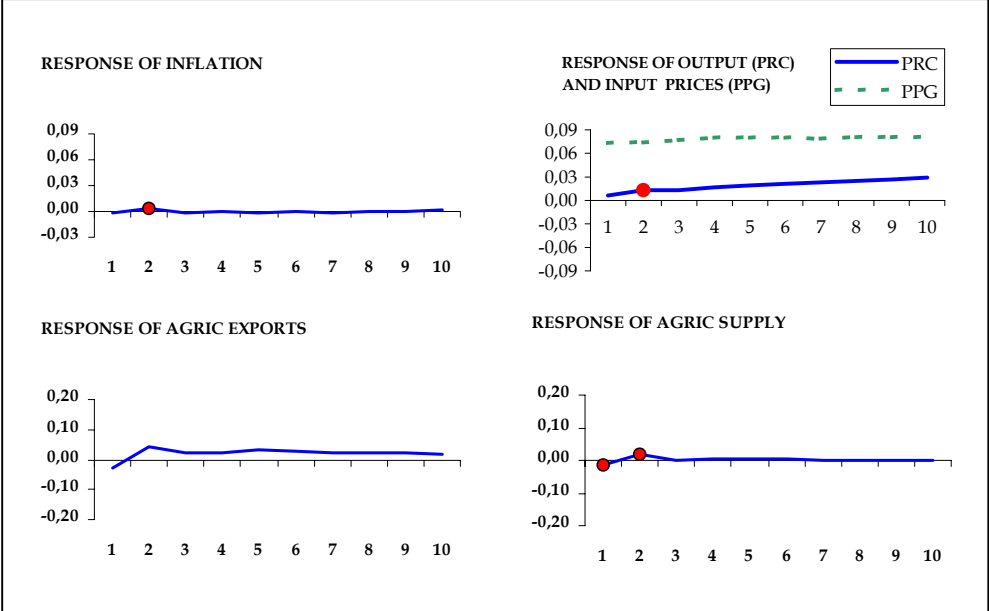
Figure 2. Responses to a shock in farm output prices



Responses to a shock in the farm input price are shown in Figure 3. Some interesting results are found. The first one is that responses of farm output prices are of lower magnitude than those of input prices. Moreover, the response of output prices is only significant two years after the initial shock. It seems that public authorities increases intervention prices as a consequence of increasing production costs and that it takes another season to producers to adapt to the new situation. In any case, and taking into account all the results found in this paper on price transmission, we can conclude that a cost-push

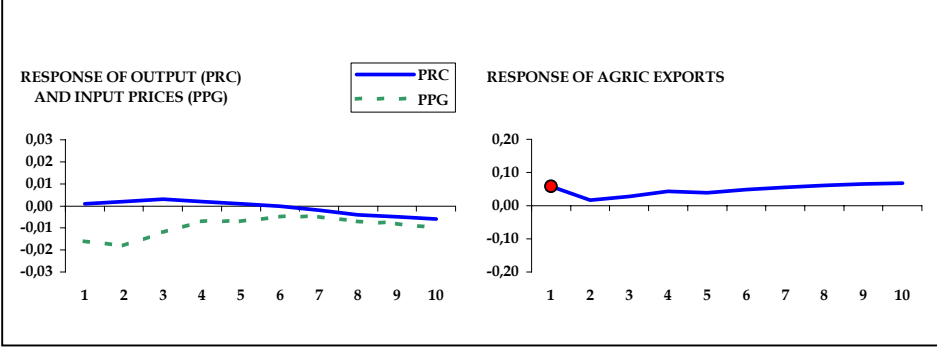
transmission mechanism prevails in the Tunisian agro-food sector. This result can be also confirmed by the significant response of inflation to increasing input prices. Finally, a positive shock in production costs reduces agricultural output the following year, which also would contribute to explain the significant increase of output prices during the second year after the shock. However, the effect on agricultural exports is not significant.

Figure 3. Responses to a shock in farm input prices



A positive shock in the exchange rate generates an immediate positive reaction in agricultural exports (Figure 4), confirming the idea we have mentioned above about the exchange rate as an important determinant of agricultural exports competitiveness. On the other hand, the effect on the rest of variables is negligible.

Figure 4. Responses to a shock in the exchange rate



5. Concluding remarks

The aim of this paper is to apply recent developments in the econometric analysis of time series to the study of relationships between macroeconomic variables and the agricultural sector in Tunisia. Results from this study suggest a number of points. The first one is that it is interesting to distinguish between long-run and short-run analyses. Long-run analysis is usually associated with structural relationships and it is in this context that theoretical restrictions have to be tested. Short-run analysis is also important for policy analysis as it gives an idea of the magnitude and time path of the reactions of economic variables to deviations from long-run relationships. However, the short-run responses of variables to shocks have to be calculated with the aid of theoretically based long-run economic restrictions.

The study has shown that changes in agricultural variables have no significant effects on macroeconomic variables. Only shocks in agricultural prices have an effect on inflation. The main source of responses of the agricultural sector (mainly agricultural output and exports) is changes in the monetary policy and, more precisely, on money supply, which is consistent with how monetary policy is instrumented in Tunisia.

Agricultural prices responses to macroeconomic shocks are not very significant as an indication of the degree of government intervention in Tunisia. In the case of agricultural exports, responses are larger if they are generated by changes in the exchange rate than if they are generated by changes in output prices, which is an indication that macroeconomic variables have to take into account when designing the agricultural policy. In the same context, the agricultural supply is quite inelastic but it reacts more to changes in the capital cost than to changes in input or output prices. To conclude, it has to be said that results presented in this paper depend on the variables and sample period chosen. Further analysis, including other variables and an extended sample period, could be conducted in the future.

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APPENDIX

Variable	Symbol	Source	Units
Exchange rate	ER	International Monetary Fund (IMF)	Tunisian Dinars/ US\$
Interest rate	R	Banque Centrale de Tunisie.	Percentage
Money Supply	M	International Monetary Fund (IMF)	Million dinars
Consumer Price Index	P	International Monetary Fund (IMF)	Index (1990 = 100)
Gross Domestic Product	GDP	International Monetary Fund (IMF)	1990 Million dinars
Rate of Commercial Openness	RCO	International Monetary Fund (IMF)	Percentage
Farm output prices	PP	Institut National de la Statistique. Tunisia.	Index (1990 = 100)
Farm input prices	IP	Institut national de la Statistique. Tunisia.	Index (1990 = 100)
Agricultural exports	AX	FAO.	1990 Million dinars
Agricultural Output	AP	Institut national de la Statistique. Tunisie.	1990 Million dinars