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MACROECONOMICS AND AGRICULTURE IN TUNISIA

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MACROECONOMICS AND AGRICULTURE IN TUNISIA

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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. The 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. The 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, Africa, dynamic relationships

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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. The 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. The results indicate that changes in macroeconomic variables have an effect on the agricultural sector but the reverse effect does not hold.

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 the economy started to recover and exports dramatically increased, being one of the main contributors to economic development. 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 the Tunisian economy. It generates around 14% of the total GDP, employs 22% of the total labour force and agro-food exports represent around 15% of total exports, although they still depend to a great extent on weather conditions. Moreover, since 1986, the agricultural sector has been 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 the trade balance. The agricultural policy was, then, oriented in 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, fruit, vegetables, etc) to finance the agricultural trade deficit.

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In many cases, the 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, more dependent on international markets and, thus, 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 data limitations into account 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. As far as we know, 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 and exchange rates (Arfa, 1994; Ben Said, 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 review of the literature 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 the 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 the exchange rate and on 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 not only the exchange rate, but income, the interest rate, as well as the usual agricultural variables, are considered 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

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3 main issue is whether levels of agricultural and non-agricultural prices respond
4 proportionally to changes in the level of money supply in the long run, and whether
5 money is neutral in the short run. The question of money neutrality in the agricultural
6 sector, as well as the speed of price adjustments, have been considered of central
7 importance for policy analysis (Bordo, 1980; Tweeten, 1980, Bessler and Babula, 1987;
8 Devadoss and Meyers, 1987; Taylor and Spriggs, 1989; Robertson and Orden, 1990;
9 Larue and Babula, 1994; Dorfman and Lastrapes, 1996; Loizou et al., 1997; Kargbo,
10 2000, among others). Bordo (1980) argues that agricultural commodities tend to be
11 more highly standardised and therefore exhibit lower transaction costs than
12 manufactured goods. Consequently, agriculture is characterised to have rather short
13 term contracts, which lead a faster response to a monetary shock. Alternatively,
14 Tweeten (1980) argues that price shocks stemming in oligopolistic non-agricultural
15 sector and accommodated by expansionary monetary policy, cause inflation and place
16 agricultural in a price-cost squeeze.
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29 Results from most of the above-mentioned studies differ substantially from each
30 other and, in many cases, are even contradictory. There are alternative explanations for
31 such differences: the samples are not homogeneous, the number of variables included
32 differs as well as their treatment as endogenous or exogenous, and they use different
33 methodological approaches. Moreover, as Ardeni and Freebairn (2002) pointed out,
34 many studies lack an appropriate treatment of the time series properties of data implying
35 misleading results especially in the case of earlier research. However, there seems to be
36 a consensus on the fact that models analysing macroeconomic linkages to the
37 agricultural sector should include the more relevant macroeconomic variables of the
38 country being analysed and should treat them as endogenous (Devadoss et al., 1987;
39 Taylor and Spriggs, 1989; Denbaly and Torgerson, 1991; Thraen et al., 1992; Devadoss
40 and Chaudhary, 1994; In and Mount, 1994; Ben Kaabia and Gil, 2000; Ivanova et al.,
41 2003, among others). Partly for this reason, most of the analyses on this topic have
42 recently been conducted using Vector Autoregression (VAR) models. This is also the
43 methodological approach we have followed in this paper although adapted to take data
44 limitations and their stochastic properties into account.
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57 The paper is organized as follows. The data used in this study, their stochastic
58 characteristics and the methodological approach are presented in Section 2. Long-run
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3 equilibrium relationships are analysed in Section 3. The short-run dynamics are
4 considered in Section 4. Finally, some concluding remarks are outlined.
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9 10 **2. Data and methodological approach**

11 Since Sims' (1986) seminal paper, VAR models have been one of the most widely
12 used tools to analyse the dynamic relationships between macroeconomic and
13 agricultural variables. In VAR models, all variables are considered endogenous and no
14 zero/one restrictions are imposed on the variables in the system. Moreover, it is possible
15 to calculate the short-run responses to a shock in one variable in the system from any
16 other variable, by offering a convenient way to characterise data without involving
17 economic theory to restrict the dynamic relationships among variables. Cooley and
18 LeRoy (1985), among others, have criticised the usefulness of such an atheoretical
19 approach for policy analysis. To overcome this problem, "Structural" VAR (SVAR)
20 models have been used (Bernanke, 1986, Sims, 1986, and Blanchard and Quah, 1989)
21 which allow the researcher to specify and test restrictions based on economic theory
22 prior to calculating the impulse response functions (Orden and Fackler, 1989).
23 However, the economic theory driving the restrictions is "weak"; although the
24 identifying restrictions imposed are consistent with economic theory, they have not
25 been derived from fully specified economic models (see Cooley and Dwyer, 1998). To
26 resolve this dilemma, Pesaran and Shin (1998) have proposed the use of generalised
27 impulse response functions to compute short-run dynamics for a set of variables, which,
28 unlike the traditional impulse response analysis, is invariant with respect to the ordering
29 of those variables.
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46 Finally, recent developments in time series analysis have modified the
47 econometric framework for analysing the relationships between macroeconomic
48 variables and the agricultural sector. The concepts of non-stationarity and cointegration
49 have become very popular and have to be explicitly tested to properly specify an
50 econometric model. In this new context, Johansen (1988) and Johansen and Juselius
51 (1990, 1992 and 1994) provide an interesting methodology that allows the researcher to
52 distinguish between the short and the long run. On the one hand, it is possible to
53 identify the long-run structural relationships among a set of variables and how variables
54 in the system adjust to deviations from such long-run equilibrium relationships. On the
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3 other hand, it is possible to calculate the impulse response functions in a similar way to
4 that in the VAR models. This distinction is useful as economic restrictions are
5 considered to be long-run in nature while it is also interesting, for policy analysis, to
6 know how the system adjusts to disequilibrium.
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11 In this paper we have followed this methodological approach although we have
12 introduced some modifications in order to adapt it to the data limitations. Availability of
13 data is a major problem for economic modelling in Tunisia. It is difficult to find a large
14 enough sample period for many economic variables. In this study, 10 variables have
15 been considered which collect the most important information related to
16 macroeconomic variables and the agricultural sector (see the Appendix for data sources
17 and units of measurement):
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24 a) Real exchange rate (ER), defined as national currency (TND) per US dollar taking
25 into account both the US and Tunisian consumer price indices¹;
26
27 b) Real money supply² (M) (money supply (M₂) divided by the consumer price
28 index);
29
30 c) Interest rate (R), defined as the one-year money market interest rate;
31
32 d) Inflation (P) expressed as the Consumer Price Index in first differences;
33
34 e) Real Gross Domestic Product (GDP);
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36 f) Real farm output prices (PP), calculated as nominal farm output prices divided by
37 the Consumer Price Index);
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39 g) Real farm input prices (IP)³, calculated as nominal farm input prices divided by
40 the Consumer Price Index);
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51 ¹ A multilateral rather than a bilateral real exchange rate would have provided better information about
52 Tunisia's competitiveness in the trade market and on agricultural exports dynamics. However, these data
53 have only been available since 1983. In any case, when comparing the evolution of both rates since the
54 information has been available, there do not seem to be any significant differences in relation to trends
55 and turning points.

56
57 ² The objective of the monetary policy in Tunisia has been to keep inflation close to that of its main
58 competitors. Traditionally, the government establishes the growth rate of the money supply (M₂) 2%
59 lower than the expected growth of the GDP. However, this objective has always been subject to revision
60 within the year.

³ In this study the "Index de pris de vente d'engrais" has been used as an aggregate input price index was
not available for the whole sample. This index comprises most of the intermediate inputs used in
agriculture and is published by the Institut National de la Statistique of Tunisia. The same index was used

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3 h) Real agricultural exports⁴ (AX), calculated as the nominal exports value divided
4 by the Consumer Price Index);
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7 i) Agricultural output (AP), calculated as the value of the Tunisian Agricultural
8 Output divided by the Consumer Price Index; and
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11 j) Rate of commercial openness (RCO) calculated by dividing the international trade
12 flows (imports + exports) by the GDP. This variable provides an indication on
13 how the Tunisian economy is inserted into the world trade.
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17 All the variables are in logarithms, except for the interest rate and inflation, which
18 are in a percentage form and are divided by one hundred to make the estimated
19 coefficients comparable with the logarithmic changes. The sample period covers annual
20 data from 1967 to 2002. Time series univariate properties have been examined by using
21 unit root tests. As in small samples such tests have limited power (Blough, 1992), two
22 alternative unit root tests developed by Elliot et al., (1996) and Ng and Perron (2001) as
23 well as the stationary test from Kwiatkowski et al. (1992) (KPSS) have been applied.
24 All the tests indicated that all the variables were I(1)⁵.
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32 Moreover, in order to check the impact of the implementation of the 1986
33 Structural Adjustment Program on the evolution of the above mentioned variables, unit
34 root test allowing for the presence of structural breaks have been performed. From the
35 seminal work by Perron (1989) on the influence of structural breaks on unit root tests,
36 much of the literature have dealt with the case in which a break occurs during one
37 period only. However, it may be more reasonable to think that breaks occur over a
38 number of periods and display smooth transition to a new level. Saikkonen and
39 Lütkepohl (2002) and Lanne et al. (2002) develop such a model which adds to the
40 deterministic term shift functions of a general nonlinear form $f_t(\Phi)$. In a model with a
41 linear trend term and shift the underlying model is given by:
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52 by Dhehibi and Lachaal (2006) to calculate the sources of agriculture productivity growth in Tunisia
53 during the period 1961-2000.

54 ⁴ Agricultural exports in Tunisia mainly refer to olive oil (40% of total agricultural exports), fish (20%);
55 dates (10%) and citrus fruit (3%). Around 60% is exported to the European Union (EU). Except for the
56 dates, the EU set a maximum amount to be imported with a lower tariff. Only in the case of the olive oil,
57 in some years total exports have exceeded the maximum amount allowed. Theoretically, the Tunisian
58 government does not directly subsidise exports. However, from the 60's to the mid 80's agricultural
59 policy was based on guaranteed prices and subsidies for agricultural inputs for selected agricultural
60 products (mainly food staples but also for some exporting products).

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

$$y_t = \mu_0 + \mu_1 t + f_t(\Phi)' \gamma + z_t \quad (1)$$

where Φ and γ are unknown parameters and z_t are residual errors. The shift function $f_t(\Phi)$ is able to characterize whether the break is abrupt and complete within one time period or more gradual. For instance, consider the simple shift function with break at time T_B :

$$f_t = d_{it} = \begin{cases} 0, & t < T_B \\ 1, & t \geq T_B \end{cases} \quad (2)$$

In this case, the structural break is represented by a simple dummy variable with shift date T_B . On the other hand, a more general shift function which allows for sharp, one-time shifts and a more gradual shift to a new level beginning at time T_B can be expressed as:

$$f_t(\Phi) = \begin{cases} 0, & t < T_B \\ 1 - \exp\{-\Phi(t - T_B + 1)\}, & t \geq T_B \end{cases} \quad (3)$$

As can be observed, in this latter case, when $\Phi \rightarrow \infty$, the shift function becomes a dummy variable. Saikkonen and Lütkepohl (2002) and Lanne et al. (2002) propose unit root tests for the model (1) in first differences which are based on estimating the deterministic term first by a generalized least squares (GLS) procedure under the unit root null hypothesis and subtracting it from the original series. Then an Augmented-Dickey-Fuller (ADF) type test is performed on the adjusted series, which also includes terms to correct for estimation errors in the parameters of the deterministic part. As in the case of the ADF statistic, the asymptotic null distribution is nonstandard. Critical values are tabulated in Lanne et al. (2002).

The user of the test has to decide on the AR order and the shift date. If the latter is known, the desired shift function may be included and the AR order may be chosen in the usual way with the help of order selection criteria, sequential tests and model checking tools. This approach is extended to a situation of unknown break date by Lanne et al. (2003), which is the approach followed in this paper. Results are shown in Table 1 and indicate that the null of unit root against the alternative of stationarity with a structural break can not be rejected. Finally, it is important to note that the endogenously determined structural breaks for each variable are different and it seems that the implementation of the Structural Adjusted Program has not generated an abrupt change in the evolution of the analyzed series.

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3 Taking into account the number of variables, the number of observations available
4 for each variable and that all variables are I(1), the methodological approach followed
5 in this paper consisted of the following steps:
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9 i) The ten-variable system is divided into two subsystems. The first one has been
10 defined by including: the real money supply, inflation, the GDP, the farm input
11 and output prices and the interest rate. Furthermore, taking into account the
12 characteristics of the Tunisian economy, we have considered the interest rate as
13 purely exogenous. The second subsystem includes the following seven variables:
14 the farm input and output prices, the agricultural exports, the agricultural
15 production, the exchange rate, the interest rate and the rate of commercial
16 openness. Within this subsystem, the interest rate and the rate of commercial
17 openness are defined as purely exogenous⁶.
18
19 ii) Under the assumption of exogeneity for certain variables, the multivariate
20 cointegration procedure developed by Pesaran et al. (2000) is used to test for
21 cointegration in both subsystems. Moreover, cointegration vectors are identified
22 as long-run meaningful economic relationships.
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24 iii) Merging the results from the two subsystems into a single system with the original
25 10 variables, impulse response functions are computed to analyse the short-run
26 dynamics and to test the exogeneity assumptions made in the first step.
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39 (Insert Table 1)
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43 3. Long-run analysis 44

45 3.1. Model specification and cointegration rank 46 47

48 All the variables in each subsystem were I(1) and, so, a Vector Error Correction
49 Model has been specified for each subsystem. The methodology developed by Pesaran
50 et al. (2000) is used to determine the cointegration rank. These authors modified the
51 Johansen (1988) procedure to explicitly allow for the introduction of exogenous
52 variables. The base-line econometric specification for multivariate cointegration is a
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59 ⁶ In a further step in the modelling process, specific tests will be carried out to test for the exogeneity of
60 the mentioned variables.

VAR(p) representation of a k -dimensional time series vector Y_t reparametrized as a Vector Error Correction Model (VECM):

$$\Delta Y_t = \mu D_t + \Gamma_1 \Delta Y_{t-1} + \dots + \Gamma_{p-1} \Delta Y_{t-p+1} - \Pi Y_{t-1} + e_t \quad (4)$$

where, Y_t is a $(k \times 1)$ column vector of variables; D_t is a vector of deterministic variables (intercepts, trend...); and 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 $niid(0, \Sigma)$.

When exogenous variables are considered, the Y_t vector can be partitioned as $Y_t = (Z_t', X_t')$, where Z_t is an $(m \times 1)$ vector of endogenous variables and X_t is an $(n \times 1)$ 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 .

Under such circumstances, Pesaran et al. (2000) show that the k -variable system defined in (4) can be decomposed to following two subsystems:⁷

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

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

“Variables cannot be exogenous per se” (Hendry, 1995). A variable can only be exogenous with respect to a set of parameters of interest. Hence, if the variables X_t are deemed to be exogenous with respect to parameters in (5), the marginal model (6) can be neglected and the conditional model (5) is complete and sufficient to sustain valid inference. Hence, the knowledge of the marginal model will not significantly improve the statistical or forecasting performance of the conditional model.

Following this line of reason, the conditional model in equation (5) is used to test for cointegration, which is equivalent to testing for the Rank (r) of matrix Π_z . So, the

⁷ Under this decomposition, variables in X_t are assumed to be weakly exogenous with respect to the cointegration space. Moreover, if the variables in Z_t do not Granger-cause X_t , then these variables are assumed to be strongly exogenous with respect to the cointegration space, that is, they would be only explained by their own past in the marginal subsystem.

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3 null hypothesis to test that there exists at least r cointegrating vectors among the
4 variables included in each subsystem can be formulated as:
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$$7 \quad H_r : \text{Rank} [\Pi_z] = r \quad r=0, \dots, m \quad (7)$$

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10 To test for the number of cointegrating vectors (r), Pesaran et al., (2000),
11 following Johansen (1988), proposed two statistics: the trace statistic (which tests
12 whether there are at least r cointegrating vectors against the maintained hypothesis; and
13 the λ_{\max} statistic (which tests whether there are r cointegrating vectors against the
14 alternative $r=r+1$). If the hypothesis of cointegration is not rejected ($0 < r < m$), Y_t is said
15 to be cointegrated in the sense that there exists a $k \times r$ matrix β such that $(\beta' Y_{t-1})$ is
16 stationary and, consequently, the cointegration relationships can be formally
17 expressed as $\Pi_z = \alpha_z \beta'$. Each column of matrix β represents a cointegrating vector,
18 whereas the rows of matrix α_z represent the adjustment coefficients that determine the
19 speed of adjustment of the $m-1$ equations to disequilibrium.
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29 The procedure outlined above has been applied to the two subsystems described in
30 the last section. However, in empirical applications, the choice of r is frequently
31 sensitive to: i) the deterministic terms included in the system (such as a constant and/or
32 a trend) and on the way in which such components interact with the error correction
33 term; and ii) the appropriate lag length to ensure that the residuals are Gaussian. In this
34 paper, both subsystems are estimated including two lags⁸ and a constant restricted to the
35 cointegration space⁹. Multivariate tests for autocorrelation (Godfrey, 1988) and
36 normality (Doornik and Hansen, 1994) have been carried out to check for model
37 statistical adequacy before applying the reduced rank tests. Results indicated that both
38 subsystems could be considered correctly specified¹⁰.
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47 Table 2 shows the results of the cointegration tests in both subsystems. As can be
48 observed, for the first subsystem (upper part of Table 2) the results of the λ -max and the
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52 ⁸ A small-sample adjusted Likelihood Ratio statistic has been used considering a maximum lag of three
53 periods.

54 ⁹ Results from unit root tests indicated that almost all the variables were non-stationary with no-zero
55 means.

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

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3 trace tests indicate that there are two cointegration vectors among the six variables
4 included while for the second subsystem (lower part of Table 2) the results differ
5 depending on the level of significance (two and three cointegration vectors for the 5 and
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9 10% levels of significance, respectively).

10
11 (Insert Table 2)

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13 Taking into account the relatively large dimension of the VECM and the small
14 sample available, the outcome of the test procedure has to be interpreted with some
15 caution. Several simulation studies show (Abadir et al., 1999; Gredenhoff and Jacobson,
16 2001; and Johansen and Juselius, 2000) that the asymptotic critical values may not be
17 very close approximations in small samples. Because of this, we have also studied the
18 roots of the companion matrix and the t-ratios of the α_z parameters from the last
19 cointegration vector (Juselius, 1995). For both subsystems, all the roots were inside the
20 unit circle, indicating that all the variables were I(1). Moreover, the eigenvalues of the
21 companion matrix show that, for both subsystems, the first four roots were close to
22 unity while the rest were quite small. In other words, we could not reject the null of two
23 and three cointegrating vectors for the first and the second subsystems, respectively.
24 Finally, all the t-statistics of the α_z parameters of the third cointegration vector for the
25 first subsystem were not significant, while in the second subsystem, some of them were
26 significant¹¹. Thus, the first subsystem has been specified with two cointegrating
27 vectors, whereas three cointegration vectors have been chosen for the second one.
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41 3.2 Long-run structural relationships

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43 The estimated β and α_z parameters are presented in Table 3, where β is presented
44 in normalized form. As can be observed, all the parameters of the long-run equilibrium
45 relationships found in each subsystem are statistically significant at the 5% level of
46 significance. Identifying economically interpretable relations is the primary aim of this
47 analysis. However, Juselius (1994) argues "the interpretation of the unrestricted
48 cointegration space is far from straightforward when there are more than one
49 cointegrating vector". Moreover, Johansen and Juselius (1994) suggest that only
50 sometimes can the unrestricted cointegrating vectors, surprisingly, be directly
51 interpreted in terms of theoretical economic relationships. Thus, some restrictions are
52 needed in order to obtain a structural representation of such relationships.
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¹¹ Results are not presented due to space limitations but are available from the authors upon request.

(Insert Table 3)

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 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 playing 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: $-\frac{\beta_{\text{IP}}}{\beta_{\text{PP}}} = \frac{\partial \text{IP}}{\partial \text{PP}} \frac{\text{PP}}{\text{IP}} = 1$

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 the data. In the first step, each single restricted relation (8)-(9) is tested for stationarity, leaving the other relations unrestricted. In other words, to test whether the 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 $H_{0i}: \beta=(H_i\phi,\omega)^{12}$. In such an expression, the restrictions to be tested are only placed in a single cointegration vector while the 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 is any 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 the main results found, is shown in Table 4. With respect to the first relationship, three different hypotheses have been tested. In the first one (H_{01}^{sys1}), it is tested whether real money is cointegrated with the interest rate, GDP and inflation, imposing income homogeneity as well. Results from the Likelihood Ratio (LR) statistic indicate that the null cannot be rejected¹³. 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 by taking an aggregate money stock into account. The third hypothesis (H_{03}^{sys1}) is similar to the first one but excluding income homogeneity. Also in this case, we fail to reject the null hypothesis. Finally, with respect to the second relationship, the 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 the results shown in the upper part of the Table 4 into account, two hypotheses have been tested. The first one (H_{05}^{sys1}) jointly tests hypotheses H_{01}^{sys1} and H_{04}^{sys1} , whereas the second tests hypotheses

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

¹³ 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. In addition, some simulation studies have shown that, in small samples, the use of the χ^2 critical values can generate considerable size distortions when testing for hypotheses on the cointegration space (Gredenhoff and Jacobsen, 1998). Garratt et al. (1999) used 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.

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4 H_{03}^{sys1} and H_{04}^{sys1} . Only in the second case do we fail to reject the null hypothesis (the LR
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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, inflation plays a significant role in the demand for money and that agricultural prices satisfy the homogeneity condition.

(Insert Table 4)

Second subsystem

In the second subsystem, taking into account the variables included and results obtained in the first subsystem concerning 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 the 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:

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

In order to test the 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 (Table 5). 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 the consistency of the data used. So, we have tested whether agricultural price homogeneity is stationary. The results from the LR test indicate that the null cannot be rejected, the same result 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, the exchange rate and the rate of commercial openness. The second one (H_{03}^{sys2}) excludes the rate of commercial openness and includes 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 rate of commercial openness 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. So, 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, hypothesis H_{09}^{sys2} jointly tests H_{01}^{sys2} and H_{05}^{sys2} . The results are shown in the middle part of Table 5. 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 5, 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.

(Insert Table 5)

Finally, Table 6 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 real money supply and the farm output prices. In the second subsystem, the three cointegrating vectors have been normalised by the farm output

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3 prices, the agricultural exports and agricultural supply, respectively. All the coefficients
4 are statistically significant and have the expected theoretical signs. The only exception
5 is the positive sign of inflation in the money demand equation. This result, following
6 Sriram (1999), would indicate that, in Tunisia, an expectation of increasing inflation
7 would drive economic agents to accumulate money stock to increase nominal income.
8 The signs associated with the rate of commercial openness in the agricultural export and
9 production equations are also interesting. In the first case, the sign is positive as
10 expected. In the second equation, the sign is negative, suggesting that a higher rate of
11 commercial openness would generate increasing imports of food products in which
12 Tunisia does not have competitive advantage (cereals, beef, vegetable oils, etc.), which
13 negatively affects domestic production.
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23 Juselius (1999) points out that "it is no longer possible to interpret a coefficient in
24 a cointegrating relation as in conventional regression context....In multivariate
25 cointegration analysis all variables are stochastic and a shock to one variable is
26 transmitted to all variables via dynamics of the system until the system has found its
27 new equilibrium position". So, the magnitude of the coefficients cannot be interpreted.
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33 On the other hand, in this type of analysis, it is also convenient to consider the
34 estimated $\alpha_{i,j}^z$ (i indicates the row and j the column) parameters as they provide valuable
35 information about the speed of adjustment of each variable towards the long-run
36 equilibrium. As the relationships between macroeconomic variables and the agricultural
37 sector are of interest for this study, let us focus on such relationships.
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43 (Insert Table 6)
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45 In relation to the first subsystem, and only considering the money demand and the
46 price equations, the first conclusion is that there seems to be a feedback relationship
47 between macroeconomic variables and the agricultural sector. In fact, any shock in the
48 money demand generates a response of input and output prices. On the other hand, any
49 change in the long-run relationship between agricultural prices affects both the income
50 (Gross Domestic Product) and inflation. In relation to the price transmission
51 mechanism, although in the long-run homogeneity holds, in the short-run the situation
52 looks different. The α_z parameters corresponding to the first cointegrating relationship
53 indicate that input prices react quicker than output prices. This result suggests a cost-
54 push transmission mechanism within the Tunisian agricultural sector, which is also
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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 does not seem to be 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 than on the agricultural production, for example, commercial agreements (most of the exported food products are sent to the European Union and are subject to contingents) or decisions made by existing exporter lobbies in the most important export goods (olive oil, dates, citrus fruit, etc.) (see also, Allaya, 1995). In other words, agricultural policy is more oriented to supporting agricultural prices and producers' and consumers' income than to encouraging trade competitiveness. Moreover, parameter α_{31} is not significant, indicating that there is no significant relationship between farm output prices and agricultural exports, which reinforces the idea of dissociation between agricultural exports and supply.

However, simply considering the magnitude of the 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.

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:

$$B_n = \Phi_1 B_{n-1} + \Phi_2 B_{n-2} + \dots + \Phi_k B_{n-p}; B_0 = I_p; B_n = 0 \quad \text{for} \quad n < 0; \Phi_1 = I + \Pi + \Gamma_1;$$

and $\Phi_i = \Gamma_i - \Gamma_{i-1}$ ($i=2, \dots, p$). Following Pesaran and Shin (1998) the scaled Generalized

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3 Impulse Response Functions (GIRF) of variable Y_i with respect to a standard error
4 shock in the j^{th} equation can be defined as:
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$$7 \text{ GIRF } (Y_{it}, Y_{jt}, h) = \frac{e_i' B_h \Sigma e_j}{\sqrt{\sigma_{jj}}}; \quad h = 0, \dots, n$$

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11 where $e_s(s=i, j)$ is the s^{th} column of the identity matrix.
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14 The GIRF are unique and do not require the prior orthogonalisation of the shocks
15 (the reordering of the variables in the system). On the other hand, the GIRF and the
16 orthogonalised IRF (Cholesky) coincide if the covariance matrix, Σ , is diagonal and $j=1$.
17 Standard deviations of impulse responses are obtained following Pesaran and Shin
18 (1998).
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23 To analyse the short-run dynamics, as mentioned in section 2, we have first
24 integrated the 10 variables into a full system including all the restrictions on the long-
25 run parameters shown in the previous table. As the price transmission long-run
26 relationship was specified in the two subsystems, only four cointegration relationships
27 were defined.
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33 Moreover, when defining the full system, two additional tests were carried out.
34 The first was to check if the variables we considered as purely exogenous (R and RCO)
35 really were. In this case, we have assumed that both variables followed a univariate
36 autoregressive process (i.e. they are not influenced by past values of the other variables
37 in the system) and we have tested for the significance of the appropriate Γ_i parameters
38 corresponding to such equations. The value of the statistic was 31.02, which was under
39 the critical value ($\chi^2_{17} = 33.41$) at the 1% level of significance). Second, all the
40 adjustment coefficients (α_z parameters), which were non-significant in Tables 3 and 4,
41 were restricted to zero. The test indicated that it was not possible to reject the null (the
42 LR statistic was 58.92, which was under the critical value of a $\chi^2(43)=66.95$, at the 1%
43 significance level). The final estimated model is given by:
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$$\beta'Y_i = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,014 & 0,000 & 0,586 \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,207 \\ 0,000 & 0,000 & 0,000 & 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ 0,000 & 0,000 & 0,000 & 0,000 & 0,871 & 0,000 & 1,000 & 0,000 & 0,074 & 0,894 & -0,277 \end{bmatrix} \begin{Bmatrix} M \\ P \\ GDP \\ PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \\ \text{Const.} \end{Bmatrix}$$

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

As can be observed, the long-run structural parameters are exactly the same than those shown in Table 6, which guarantees the consistency of the methodological approach followed here¹⁴. 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 of 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 positively affects inflation, although the effect is only significant during the second year after the shock. Moreover, increasing access to credits stimulates economic growth (the GDP increase) as well as agricultural exports. The effect on farm output prices is

¹⁴ As we were interested in analysing the full set of interrelationships between macroeconomic and agriculture variables, the estimation of a single system with the 10 variables would have provided more efficient estimates. However, as only 35 observations were available for each series, the estimation of a full system, as mentioned in Section 2, was not possible due to the lack of degrees of freedom. The methodological approach followed here allowed us to analyse all possible interrelationships without a significant loss of efficiency. Moreover, when comparison was possible, the impulse response functions derived from partial models were not significantly different than those shown in this paper.

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3 positive but non significant. The aim of the Tunisian agricultural policy is to support
4 farmers' income through reasonably high intervention prices although keeping inflation
5 controlled. Agricultural prices are not allowed to increase over expected inflation in
6 order to guarantee consumer access to basic foods. Limited increases of agricultural
7 prices do not stimulate either agricultural production or the demand for inputs. As a
8 result, the impact on input prices is not significant. Finally, following the Keynesian
9 theory, an expansive monetary policy induces exchange rate depreciation, which, leads
10 to an increase of agricultural exports.
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18 (Insert Figure 1)
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20 Figure 2 shows the responses to a shock in farm output prices of the most relevant
21 variables within the system. In general terms, agricultural variables do not have any
22 effect on macroeconomic variables and, so they are not included in the Figure. Two
23 main results are found. First, a positive increase of producer prices generates a positive
24 response of agricultural production. The response is significant for two years. Second, it
25 also generates an immediate positive response of input prices due to the increasing
26 derived input demand as a consequence of production increases. The magnitude of this
27 response is higher than in the case of output prices, which is consistent with the
28 comments in Section 3 about the α_z parameters and also with the trend followed by both
29 price series during the analysed period. El Abassi (1995) and Allaya (1995) explain this
30 trend by the intensification process that the Tunisian agriculture has suffered during the
31 last two decades and the progressive elimination of subsidies addressed to fertilizers and
32 other intermediate inputs. On the other hand, the effect on inflation is positive but not
33 significant. As mentioned above, the government has traditionally controlled
34 agricultural price increases to be compatible with the inflation rate in Tunisia. Finally,
35 the effect on agricultural exports is not significant either. Prices are not the main source
36 of competitiveness for Tunisian agricultural exports as they are mainly subject to
37 contingents. Moreover, the traditional policy used by the Tunisian Government to
38 promote exports has been via the exchange rate.
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54 (Insert Figure 2)
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56 Responses to a shock in the farm input price are shown in Figure 3. Some
57 interesting results are found. The first is that responses of farm output prices are of
58 lower magnitude than those of input prices, a result which is consistent with the
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3 empirical evidence provided by El Abassi (1995). Moreover, the response of output
4 prices is only significant two years after the initial shock. It seems that the public
5 authorities increase intervention prices as a consequence of increasing production costs
6 and that it takes another season for producers to adapt to the new situation. In any case,
7 and taking into account all the results found in this paper on price transmission, we can
8 conclude that a cost-push transmission mechanism prevails in the Tunisian agro-food
9 sector. This result can be also confirmed by the significant response of inflation to
10 increasing input prices. Finally, a positive shock in production costs reduces agricultural
11 output the following year, which would also contribute to explaining the significant
12 increase of output prices during the second year after the shock. However, the effect on
13 agricultural exports is not significant.

(Insert Figure 3)

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25 A positive shock in the exchange rate generates an immediate positive reaction in
26 agricultural exports (Figure 4), confirming the idea we have mentioned above about the
27 exchange rate as an important short-run determinant of the competitiveness of
28 agricultural exports. However, in the long-run the effect is not significant. This results is
29 outlined also in Ben Said (1994) and Allaya (1995) who concluded that after the
30 devaluation of the Tunisian dinar that took place with the implementation of the
31 Structural Adjustment Program the growth rate of Tunisian agricultural exports was
32 lower than that of the exchange rate. Finally, as can be also observed in Figure 4, the
33 effect on the rest of variables is negligible.

(Insert Figure 4)

46 47 **5. Concluding remarks**

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49 The aim of this paper is to apply recent developments in the econometric analysis
50 of time series to the study of relationships between macroeconomic variables and the
51 agricultural sector in Tunisia. Results from this study suggest a number of points. The
52 first is that it is interesting to distinguish between long-run and short-run analyses. The
53 long-run analysis is usually associated with structural relationships and it is in this
54 context that theoretical restrictions have to be tested. The short-run analysis is also
55 important for policy analysis as it gives an idea of the magnitude and time path of the
56 reactions of economic variables to deviations from long-run relationships. However, the
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3 short-run responses of variables to shocks have to be calculated with the aid of
4 theoretically based long-run economic restrictions.
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8 The study has shown that changes in agricultural variables have no significant
9 effects on macroeconomic variables. Only shocks in agricultural prices have an effect
10 on inflation. The main source of responses of the agricultural sector (mainly agricultural
11 output and exports) is changes in the monetary policy and, more precisely, in money
12 supply, which is consistent with how monetary policy is instrumented in Tunisia.
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17 Agricultural prices responses to macroeconomic shocks are not very significant as
18 an indication of the degree of government intervention in Tunisia. In the case of
19 agricultural exports, responses are larger if they are generated by changes in the
20 exchange rate than if they are generated by changes in output prices, which is an
21 indication that macroeconomic variables have to be taken into account when designing
22 agricultural policy. In the same context, agricultural supply is quite inelastic but reacts
23 more to changes in capital costs than to changes in input or output prices. To conclude,
24 it has to be said that the results presented in this paper depend on the variables and
25 sample period chosen. Further analysis, including other variables and an extended
26 sample period, could be conducted in the future.
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REFERENCES

- 1
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5 Abadir, K.M., Hadri, K. and Tzavalis, E. (1999). The Influence of VAR Dimensions on Estimator Biases. *Econometrica*, **67**, 163-181.
- 6 Allaya, M. (1995). Intercambios Agroalimentarios y Cooperación entre Europa y el Magreb. *Revista Española de Economía Agraria*, **172**, 31-49.
- 7 Ardeni, P. G. and Freebairn, J. (2002). The Macroeconomics of Agriculture. In: Gardner, B. and Rausser, G.C. (eds.). *Handbook of Agricultural Economics*. Vol. 2A. Amsterdam: North-Holland, 1455-1485.
- 8 Arfa, L. (1994). Les Exportations Agroalimentaires de la Tunisie vers la C.E. : Des Performances Inégales. *MEDIT*, **4**, 4-12.
- 9 Batten, D.S. and Belongia, M.T. (1986). Monetary Policy, Real Exchange Rates and U.S. Agricultural Export. *American Journal of Agricultural Economics*, **68**, 422-427.
- 10 Ben Kaabia, M. and Gil, J.M. (2000). Estimation and inference in cointegrated demand systems: an application to Tunisian meat consumption. *European Review of Agricultural Economics*, **28(3)**, 449-471.
- 11 Ben Said, T. (1994). Les Effets de l'Ajustement Structurel sur l'Intégration de l'Agriculture Tunisienne à l'Échange International. Crises et Transitions des Politiques Agricoles en Méditerranée. *Options Méditerranéennes. Série B-Etudes et Recherche*, **8**, 117-123.
- 12 Bernanke, B.S. (1986). Alternative Explanations of the Money-Income Correlation. *Cambridge, MA: National Bureau of Economic Research, Working Paper N°*. 1842.
- 13 Bessler, D.A. and Babula, R.A. (1987). Forecasting Wheat Export: Do Exchange Rate Matter?. *Journal of Business and Economic Statistics*, **5**, 397-406.
- 14 Blanchard, O.J. and Quah, D. (1989). The Dynamic Effects of Aggregate Demand and Supply Disturbances. *American Economic Review*, **79**, 655-673.
- 15 Blough, S.R. (1992). The Relationship between Power and Level for Generic Unit Root Tests in Finite Samples. *Journal of Applied Econometrics*, **7**, 295-308.
- 16 Bordo, M. D. (1980). The Effects of Monetary Change on Relative Commodity Prices and the Role of Long-Term Contracts. *Journal of Political Economy*, **88**, 1088-1109.
- 17 Cooley, T. G. and LeRoy, S. F. (1985). Atheoretical Macroeconomics: A Critique. *Journal of Monetary Economics*, **16**, 283-308.
- 18 Chambers, R. G. (1984). Agricultural and Financial Market Interdependence in the Short Run. *American Journal of Agricultural Economics*, **66**, 12-24.
- 19 Chambers, R. G. and Just, R. E. (1979). A Critique of Exchange Rate Treatment in Agricultural Trade Models. *American Journal of Agricultural Economics*, **61**, 249-257.
- 20 Cooley, T. F. and Dwyer, M (1998). Business cycle Analysis Without much Theory: A look at Structural VARs. *Journal of Econometrics*, **53**, 57-88.
- 21 Denbaly, M. and Torgerson, D. (1991). Macroeconomic Determinants of Relative Wheat Prices: Integrating the Short Run and Long Run. *Journal of Agricultural Economics Research*, **44(2)**, 27-35.
- 22 Devadoss, S. and Meyers, W. H. (1987). Relative Prices and Money: Further Results for the United States, *American Journal of Agricultural Economics*, **69**, 838-842.
- 23 Devadoss, S., Meyers, W. H. and Starleaf, D. R. (1987). Macrolinkages between the Farm and no Farm Sectors and the Impact of Monetary Policy Decisions. CARD Monograph 87-M1. Center for Agricultural and Rural Development. IOWA State University.
- 24 Devadoss, S. and Chaudhary, S. (1994). Effects of fiscal policy on US agriculture. *Applied Economics*, **26**, 991-997.
- 25 Dhehibi, B. and Lachaal, L. (2006). Empirical evidence of productivity growth in Tunisian agriculture: 1961-2000. *African Development Review*, **18**, 248-257.
- 26 Doornik, J. and Hansen, H. (1994). A Practical Test for Univariate and Multivariate Normality. Discussion paper, Nuffield College, Oxford.
- 27 Dorfman, J. H. and Lastrapes, W. D. (1996). The Dynamic Responses of Crop and Livestock Prices to Money-Supply Shock: A Bayesian Analysis Using Long-Run Identifying Restrictions. *American Journal of Agricultural Economics*, **78**, 530-441.
- 28 El Abassi, M. (1995). Ouverture et Compétitive de l'Agriculture Tunisienne. La Nouvelle Conjoncture Internationale et les Flux d'Échanges avec l'Europe. Le Cas des Produits Agricoles et Agroalimentaires. Les Agricultures Maghrébines à l'Aube de l'An 2000. *Options Méditerranéennes. Série B-Etudes et Recherches*, **14**, 337-347.
- 29 Elliott, G., Rothenberg, T. J. and Stock, J. H. (1996). Efficient Tests for an Autoregressive Unit Root, *Econometrica*, **64**, 813-836.
- 30 Garrat, A., Lee, K., Pesaran, M.H. and Shin, Y. (1999). A long Run structural Macroeconomic Model of the UK. *DAE Working Paper N° 9812*, University of Cambridge.
- 31 Godfrey, L.G. (1988). Misspecification Test in Econometrics. Cambridge University Press.
- 32 Gredenhoff, M. and Jacobson, T. (2001). Bootstrap Testing Linear Restrictions on Cointegrating Vectors. *Journal of Business and Economic Statistics*, **19(1)**, 63-72.
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3 Hendry, D. F. (1995). *Dynamic Econometrics*. Oxford: Oxford University Press.
- 4 In, F. and Mount, T. (1994). *Dynamic Macroeconomic Linkages to the Agricultural Sector*. Athenaeum Press
5 Ltd, Newcastle upon Tyne.
- 6 Ivanova, N., Dawson, P. and Lingard, J. (2003). Macroeconomic impacts on Bulgarian agriculture during
7 transition. *Applied Economics*, **35**, 817-823.
- 8 Johansen, S. (1988). *Statistics Analysis of Cointegration Vectors*. *Journal of Economic Dynamics and Control*, **12**,
9 231-254.
- 10 Johansen, S. and Juselius, K. (1990). Maximum Likelihood Estimation and Inference on Cointegration-with
11 Applications to the Demand for Money. *Oxford Bulletin of Economics and Statistics*, **52**, 169-210.
- 12 —(1992). Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and the UIP for UK.
13 *Journal of Econometrics*, **53**, 211-244.
- 14 —(1994). Identification of the Long-Run and the Short-Run Structure: An Application to the ISLM Model. *Journal*
15 *of Econometrics*, **63**, 7-36.
- 16 Juselius, K. (1994). On the Duality between Long-Run Relations and Common Trends in the I(1) Versus I(2) Model:
17 An application to aggregate money holdings. *Economic Reviews*, **13**(2), 151-178.
- 18 —(1995). Do Purchasing Power Parity and Uncovered Interest Rate Parity Hold in the Long Run? An Example
19 of Likelihood Inference in a Multivariate Time-Series Model. *Journal of Econometrics*, **69**, 211-240.
- 20 —(1998). A structured VAR for Denmark Under Changing Monetary Regimes. *Journal of Business and*
21 *Economic Statistics*, **16**(4), 400-411.
- 22 —(1999). Models and relations in Economics and Econometrics. *Journal of Economic Methodology*, **6**(2), 259-
23 290.
- 24 Kargbo, J. (2000). Impacts of monetary and macroeconomic factors on food prices in eastern and southern
25 Africa. *Applied Economics*, **32**, 1373-1389.
- 26 Kwiatkowski, D., Phillips, P., Schmidt, P. and Shin, Y. (1992). Testing the Null Hypothesis of Stationarity Against
27 the Alternative of Unit Root. *Journal of Econometrics*, **54**, 159-178.
- 28 Lanne, M., Lütkepohl and Saikkonen P. (2002). Comparison of unit root tests for time series with level shifts.
29 *Journal of Time Series Analysis*, **23**, 667- 685.
- 30 Lanne, M., Saikkonen, P. and Lütkepohl, H. (2003). Test procedures for unit roots in time series with level shifts
31 at unknown time. *Oxford Bulletin of Economics and Statistics*, **65**, 91-115.
- 32 Larue, B. and Babula, R. A. (1994). Evolving Dynamic Relationships between the Money Supply and Food-Based
33 Prices in Canada and the United States. *Canadian Journal of Agricultural Economics*, **42**, 159-176.
- 34 Loizou, E., Mattas, K. and Pagoulatos, A. (1997). Macro-monetary effects on agricultural prices: the case of Greek
35 agriculture. *Applied Economics Letters*, **4**, 397-401.
- 36 Longmire, J. and Morey, A. (1983). Strong Dollar Dampens Demand for US Farm Exports. *FAER Report n°*
37 *123*, USDA, Washington DC.
- 38 Lütkepohl, H. (1993). *Introduction to Multiple Time Series*, Springer Verlag, Berlin.
- 39 Orden, D. and Fackler, P.L. (1989). Identifying Monetary Impacts on Agricultural Prices in VAR Models. *American*
40 *Journal of Agricultural Economics*, **71**, 495-502.
- 41 Ng, S. and Perron, P. (2001). Lag Length Selection and the Construction of Unit Root Tests with Good Size and
42 Power, *Econometrica*, **69**, 1519–1554.
- 43 Pesaran, M.H. and Shin, Y. (1996). Cointegration and speed convergence to equilibrium. *Journal of*
44 *Econometrics*, **71**, 117-143.
- 45 Pesaran, M.H. and Shin, Y. (1998). Generalised Impulse Response Analysis in Lineal Multivariate Models.
46 *Economics Letters*, **58**, 17-29.
- 47 Pesaran, M.H., Shin, Y. and Smith, R.J. (2000). Structural Analysis of Vector Error Correction Models with
48 Exogenous I(1) Variables. *Journal of Econometrics*, **97**(2), 293-343.
- 49 Reimers, H.E. (1992). Comparisons of Tests for Multivariate Cointegration. *Statistical Papers*, **33**, 335-359.
- 50 Robertson, J. C. and Orden, D. (1990). Money Impacts on Prices in the Short and Long Run: Some Evidence from
51 New Zealand. *American Journal of Agricultural Economics*, **72**, 160-171.
- 52 Saikkonen, P. and Lütkepohl, H. (2002). Testing for a unit root in a time series with a level shift at unknown
53 time. *Econometric Theory*, **18**, 313-348.
- 54 Schuh, G. E. (1974). The Exchange Rate and U.S. Agricultural. *American Journal of Agricultural Economics*,
55 **56**, 1-13.
- 56 Sims, C. (1986). Are Forecasting Models Usable for Policy Analysis?. *Federal Research Bank of Minneapolis .*
57 *Quarterly Review*, **10**, 2-16.
- 58 Sriram, S. S. (1999). Demand for M2 in an Emerging-Market Economy: An Error-Correction Model for
59 Malaysia. *Working Paper WP/99/173*. International Monetary Fund. Washington.
- 60 Taylor, J. S. and Spriggs (1989). Effects of the Monetary Macroeconomy on Canadian Agricultural Price. *Canadian*
Journal of Agricultural Economics, **22**, 278-289.

- 1
2
3 Thraen, C. S, Hwang, T. C. and Larson, D. W. (1992). Linking of U.S. Monetary Policy and Exchange Rates to
4 World Soybean Markets. *Agricultural Economics*, **6**, 365-384.
5 Tweeten, L. (1980). An Economic Investigation of Inflation Passthrough to the Farm Sector. *Western Journal of*
6 *Agricultural Economics*, **5**, 89-106.
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For Peer Review

APPENDIX

Variable	Symbol	Source	Units
Exchange rate	ER	International Monetary Fund (IMF)	Tunisian Dinars per US dollar
Interest rate	R	Statistiques financières. Banque Centrale de Tunisie.	Percentage
Money Supply	M	International Monetary Fund (IMF)	Million dinars
Consumer Price Index	P	International Monetary Fund (IMF)	Index (Basis 100 = 1990)
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. Ministère du développement économique. Tunisie.	Index (Basis 100 = 1990)
Farm input prices	IP	Institut National de la Statistique. Ministère du développement économique. Tunisie.	Index (Basis 100 = 1990)
Agricultural exports	AX	FAO.	1990 Million dinars
Agricultural Output	AP	Institut National de la Statistique. Ministère du développement économique. Tunisie.	1990 Million dinars

Table 1. Results from Lanne et al. (2003) unit root tests with structural breaks.

Variable	Break point (T_B)	With trend	Break point (T_B)	Whithout trend
ER	1981	-1.82	1980	-0.33
R	1981	-1.58	1981	-1.50
M	1974	-1.81	1974	-1.56
P	1996	-2.12	1980	-1.85
GDP	1982	-2.68	1982	-2.12
RCO	1974	-1.44	1976	-2.42
PP	1982	-1.28	1982	-2.17
IP	1986	-2.59	1986	-2.56
AX	1972	-1.77	1976	-2.02
AP	1971	-2.21	1995	-0.91
CV (5%) ^a		-3.15		-2.99

^a Critical values are from Lanne et al. (2002) (Table 2, T=50).

Table 2. 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 definitions of the variables

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

Table 3. Estimated β and α_z parameters for both subsystems

β parameters ¹						
First subsystem (r=2)			Second subsystem (r=3)			
Variable	Vector 1	Vector 2	Variable	Vector 1	Vector 2	Vector 3
M*	1,000	-0,130	PP*	-2,471	-1,112	1,000
P*	-5,157	-0,764	IP*	-1,468	-3,360	-0,079
GDP*	-1,366	0,101	AX*	1,000	-0,533	0,099
PP*	-0,381	1,000	AP*	-1,068	1,000	-0,312
IP*	-0,283	-0,052	ER*	1,011	4,125	0,029
R*	0,017	-0,006	R*	0,079	0,026	0,015
Constant*	8,916	-0,231	RCO*	-1,210	-3,552	-0,105
			Constant*	2,188	-8,913	1,961
α_z parameters						
DM	-0,166	0,762	DPP	0,119	-0,015	-0,137
DP	0,175	-0,131	DIP	0,374	0,045	-0,666
DGDP	0,213	0,541	DAX	-0,431	0,100	-2,183
DPP	-0,068	-0,015	DAP	0,218	0,016	1,253
DIP	-0,292	-0,159	DER	-0,152	-0,018	-0,455

¹ An * indicates that the parameter is significant at the 5% level of significance (critical values are 5.99 and 7.81 for the first and second subsystems, respectively)

Table 4. Hypothesis restriction 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.

Table 5. Hypothesis restriction tests on the cointegration vectors in the second subsystem^a

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

Table 6. Estimated β and α_z matrices under long-run identification for both subsystems
First subsystem

$$\beta' = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,014 & 0,586 \\ & (0,413) & (0,026) & & & (0,007) & (0,231) \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,207 \\ & & & & & & (0,471) \end{bmatrix} \times \begin{bmatrix} M \\ P \\ GDP \\ PP \\ IP \\ R \\ Const. \end{bmatrix}$$

$$\alpha_z = \begin{bmatrix} -0,123 & -0,014 \\ (-1,827) & (-0,522) \\ 0,156 & 0,034 \\ (4,343) & (3,752) \\ 0,204 & 0,066 \\ (3,034) & (3,883) \\ 0,096 & -0,023 \\ (1,945) & (-2,787) \\ 0,314 & 0,076 \\ (1,978) & (2,691) \end{bmatrix}$$

Likelihood Ratio Statistic= 11,55
Critical value (1%) = 15,09

Second subsystem

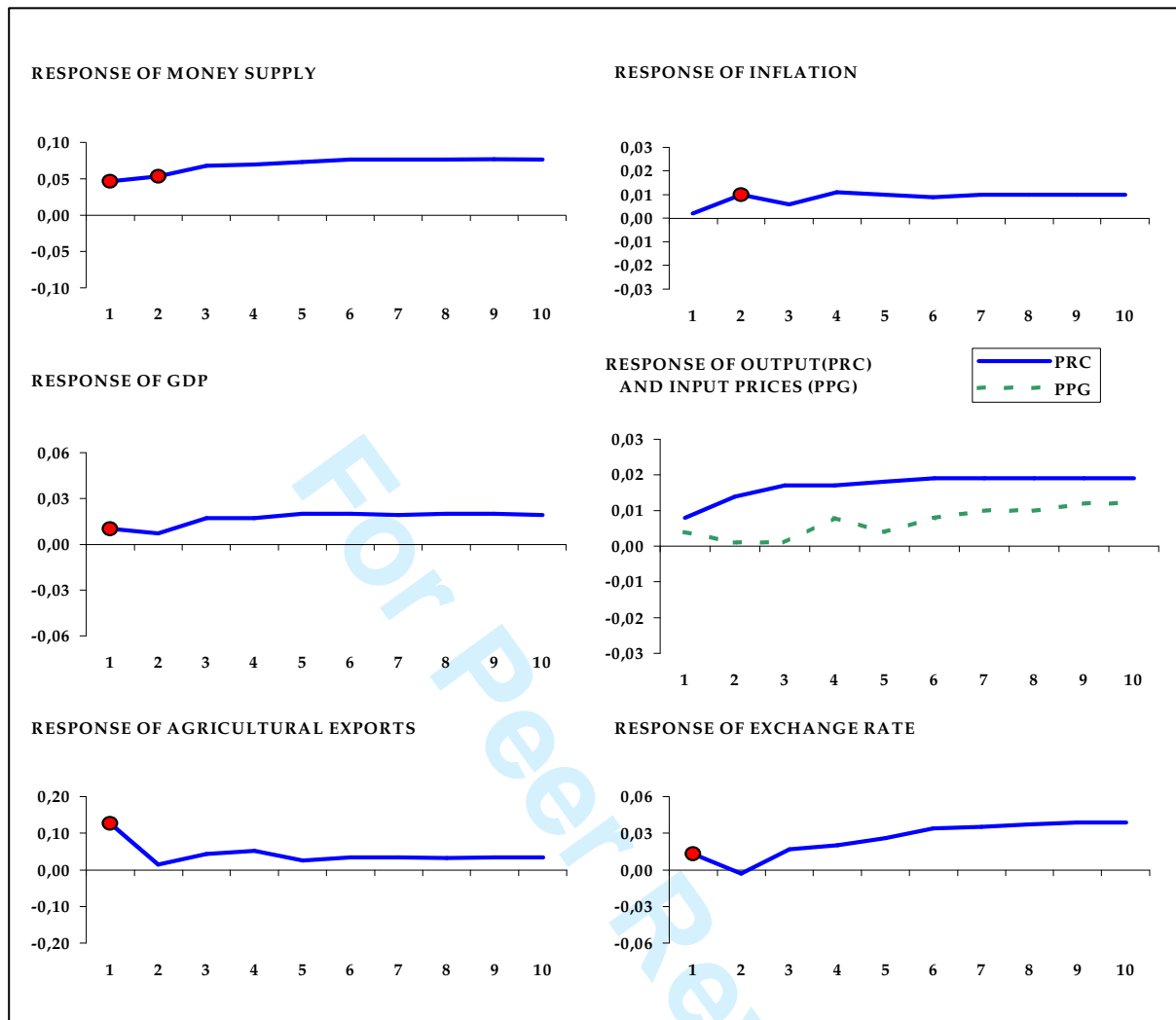
$$\beta' = \begin{bmatrix} 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,269 \\ & & & & & & & (0,058) \\ 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ (0,384) & & & & (0,083) & & (0,194) & (0,823) \\ 0,000 & 0,871 & 0,000 & 1,000 & 0,000 & 0,074 & 0,894 & -0,277 \\ & (0,089) & & & & (0,080) & (0,178) & (0,862) \end{bmatrix} \times \begin{bmatrix} PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \\ Const. \end{bmatrix}$$

$$\alpha_z = \begin{bmatrix} -0,146 & 0,079 & -0,049 \\ (-2,279) & (2,781) & (-2,083) \\ 0,260 & 0,109 & -0,361 \\ (4,014) & (4,328) & (-3,871) \\ -0,217 & -0,363 & 0,400 \\ (-0,832) & (-3,898) & (0,786) \\ 0,145 & 0,260 & -0,128 \\ (2,356) & (1,323) & (-1,932) \\ -0,048 & -0,028 & 0,003 \\ (-0,723) & (-1,198) & (0,873) \end{bmatrix}$$

Likelihood Ratio Statistic = 15,23
Critical Value (1%) = 16,81

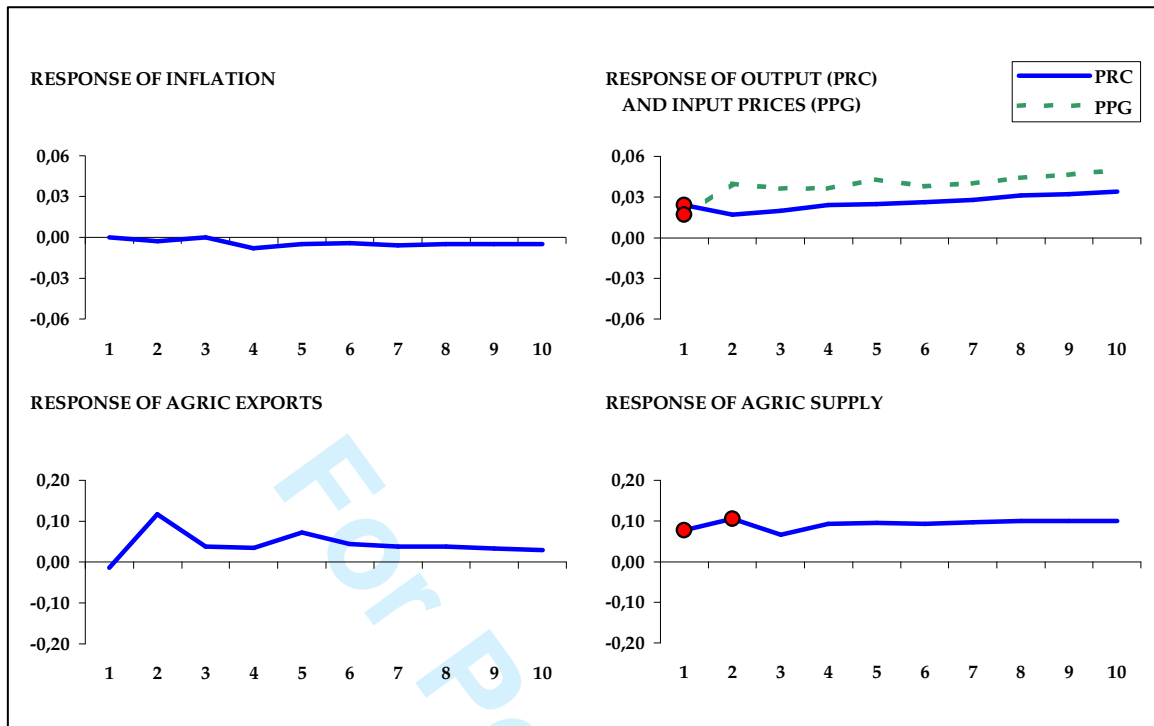
Note: Values in parentheses correspond to standard deviations, in the case of the β parameters, and to t-ratios, in the case of the α_z parameters

Figure 1. Responses to a shock in the money supply



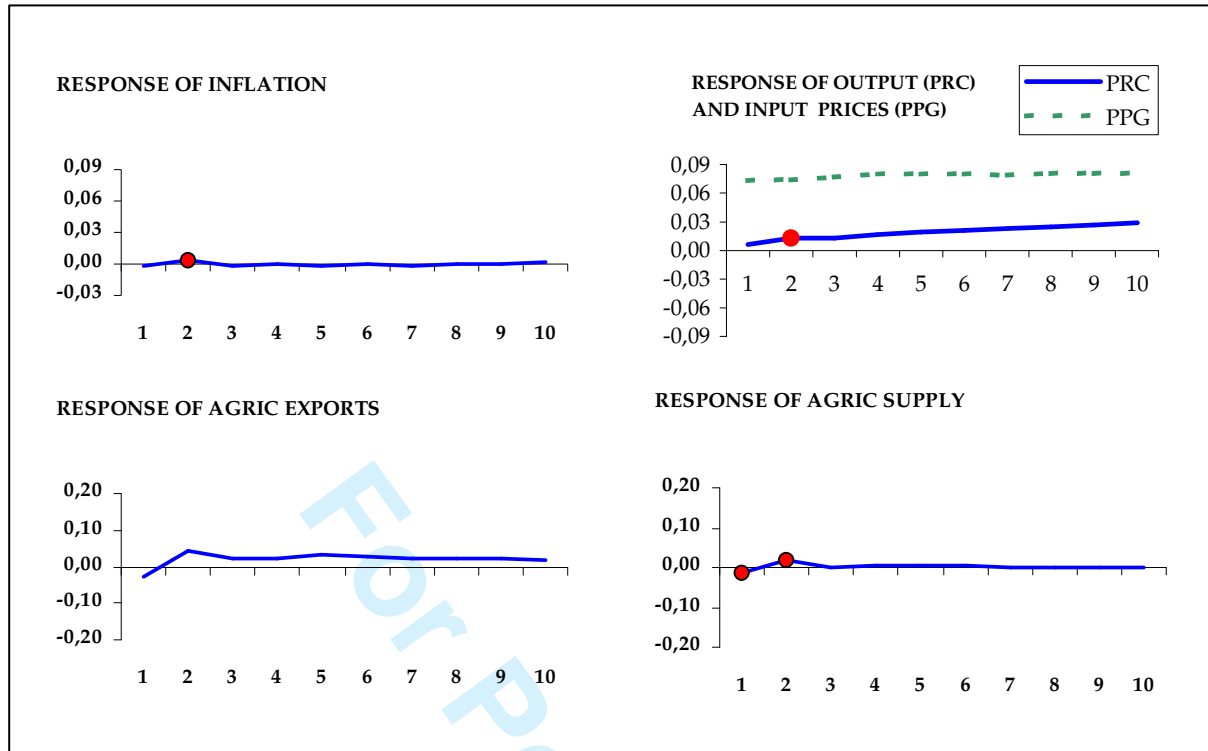
Note: Significant responses at the 5% level of significance are marked with a circle

Figure 2. Responses to a shock in farm output prices



Note: Significant responses at the 5% level of significance are marked with a circle

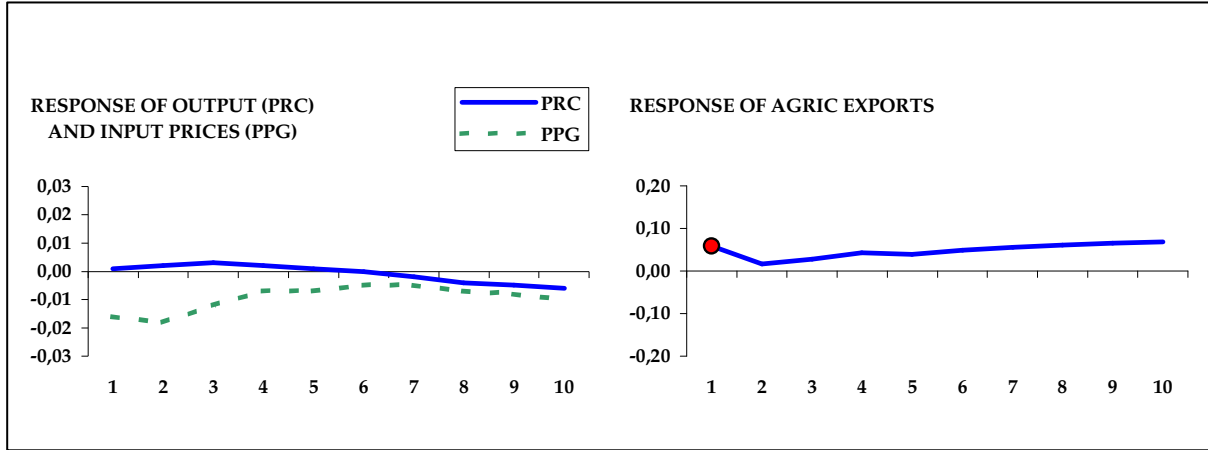
Figure 3. Responses to a shock in farm input prices



Note: Significant responses at the 5% level of significance are marked with a circle

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Figure 4. Responses to a shock in the exchange rate



Note: Significant responses at the 5% level of significance are marked with a circle