The effect of long-run identification on impulse-response functions: An application to the relationship between macroeconomics and agriculture in Tunisia

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Abstract
The objective of this paper is to analyse some of the most relevant relationships among macroeconomic variables and the agricultural sector in Tunisia. Three alternative models are specified and estimated: a VAR in levels, an unrestricted Vector Error Correction Model (VECM), and a Restricted VECM in which long-run relationships among the relevant variables are identified. In all models short-run dynamics are analysed through the use of Generalised Impulse Response Functions. Results indicate that alternative model specifications generate different short-run dynamics. Long-run identification seems to be a necessary condition for obtaining consistent economic results.

Key words: *Cointegration; Identification; Impulse response function.*

Introduction

Since the mid seventies many empirical and theoretical studies have demonstrated that the macroeconomic environment exerts a substantial influence on agricultural sector[1]. Since the Sims' (1986) seminal paper, VAR models have been one of the most widely used analysis tools in this particular field[2]. One of the advantages of this methodology is that all variables are treated as endogenous and no zero/one restrictions are imposed on the variables in the system. In addition, VAR models are particularly suitable to studying the time-path response of agricultural variables to unexpected shocks using impulse response functions (IRF) and the decompositions of the forecast error variances (Bessler and Babula, 1987; Chambers, 1984; Orden and Fackler, 1989). However, as often recognised in the literature, this technique is subject to criticism. The main complaint being that the estimated IRF requires a previous orthogonalisation (identification) of the shocks in order to give an economic interpretation for the source of the shock. This identification can be achieved by using either a Cholesky or a structural decomposition. The Cholesky decomposition implicitly assumes a recursive contemporaneous structure (the so-called Wald causal chain). The resulting IRF will not be unique, and it will depend on the ordering of the variables in the model. In order to avoid an arbitrary ordering, Bernanke (1986), Blanchard and Quah (1989), proposed the so-called structural VAR model, where appropriate restrictions are imposed on the system variables.

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These restrictions are derived from economic theory. However, the economic theory driving the restrictions is "weak"; although the identifying restrictions imposed are consistent with economic theory, they have not been derived from fully specified economic models (see Cooley and Dwyer, 1998). To resolve this dilemma, Pesaran and Shin (1998) have proposed the use of generalised impulse response functions to compute short-run dynamics for a set of variables, which, unlike the traditional impulse response analysis, is invariant with respect to the ordering of those variables. *On the other hand, if certain variables within the system are cointegrated, then there exist some long-run structural relationships among them. Earlier papers using VAR models did not consider the stochastic properties of the data. The IRF were directly calculated from a model in levels or, at least, in first differences, while ignoring the underlying long-run structure. Recently, more attention has been paid to identify long-run relationships but only in a few cases have the short-run dynamics been addressed. Naka and Tufte (1997) examined the consequences of ignoring cointegration on impulse response functions. In their paper, two models were considered: a VAR model in levels with one lag and an equivalent Vector Error Correction model (VECM). A Monte Carlo experiment suggested that both estimation methods would yield reasonably similar IRF at short horizons. However at longer horizons, those from VAR in levels diverge from their true value. This finding was consistent with those presented in Engle and Yoo (1989); and Lin and Tsay (1996) on forecast accuracy.*

The objective of this paper is twofold. First, we want to analyse the relationship between certain macroeconomic variables and the agricultural sector in Tunisia. This may well be one of the first studies of this kind in this country. Second, we aim to provide more insight about the sensitivity of IRF to various model specifications extending the mentioned work by Naka and Tufte (1997) as we are considering also a model in which long-run identification is achieved. Results will show that different model specifications generate different IRF. These differences might have important consequences for economic policy.

Data and non-stationarity

Availability of data is a major problem for economic modelling in Tunisia. For this reason we have restricted our analysis only to a few variables for which homogeneous and long enough information has been found. In any case, as the main objective is to assess the effects of alternative model specifications on the impulse-response functions, our analysis is therefore restricted to evaluate the effects of money supply and the exchange rate on agricultural prices and exports, two issues which have been the subject of research interest among agricultural economists.

Four variables have been considered: 1) the real exchange rate (RER)[3], defined as national currency (TND) per US dollar taking into account both the US and Tunisian consumer price indices; 2) Real money supply[4] (RM) (money supply divided by the consumer price index); 3) Real farm output prices (RAP) (calculated as nominal prices divided by the consumer price index); and 4) Real agricultural exports[5] (RAX) (exports value divided by the consumer price index). All variables are expressed in logarithms. Annual data from 1963 to 1998 are used. Figure 1 shows the evolution of the four variables.

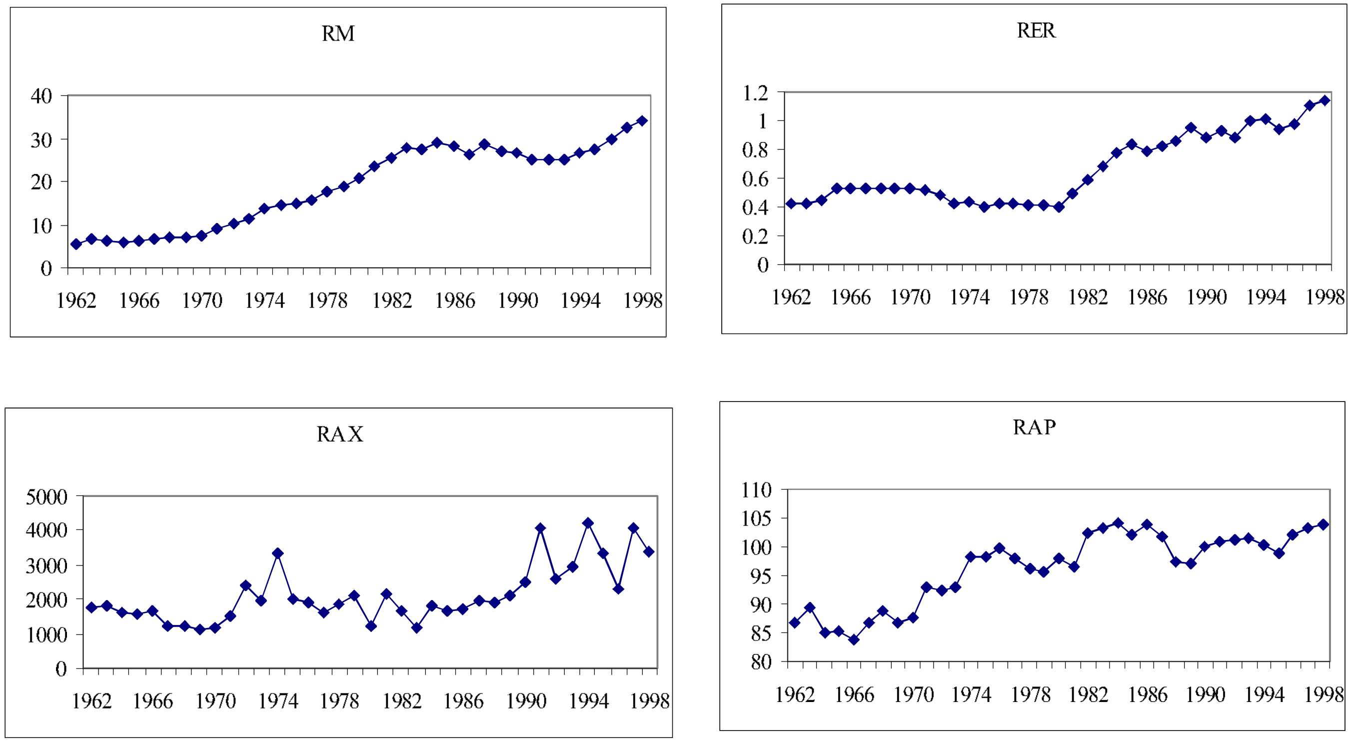

Figure 1. Evolution of the agricultural and macroeconomic variables included in the system

Taking into account the methodological approach followed in this paper, the first step in our analysis has been to test the order of integration of each series. When the number of observations is low, unit root tests have little power. For this reason we have examined the results from two different tests: the Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979, 1981), which test the null of unit root, and KPSS (Kwiatkowski et al., 1992), which tests the null of stationarity. Both tests indicated that the four variables were clearly I(1) [6].

Cointegration analysis and short-run dynamics

The cointegration analysis has been conducted using the Johansen approach. Johansen's (1988) procedure starts with the following reformulation of a VAR(k) model into a Vector Error Correction Model (VECM):

$$\Delta Z_t = \Pi Z_{t-1} + \delta D_t + \sum_{i=1}^{k-1} \Psi_i \Delta Z_{t-i} + \varepsilon_t \tag{1}$$

where Z_t is a p×1 vector of endogenous variables; ψ_i, i= 1, 2… are (p×p) matrices of short-run parameters; Π is a (p×p) matrix of long-run parameters; D_t is a vector of deterministic terms (a constant, a linear trend, seasonal dummies, intervention dummies, etc.); and ε_t is a vector of errors that are assumed to be independently and identically Gaussian distributed, such that $E(\varepsilon_t \varepsilon_t') = \Sigma$ for all t, where $\Sigma = \{\sigma_{ij}, i, j = 1,2,..., p\}$ is an (p×p) positive definite matrix.

In the I(1) system Z_t is said to be cointegrated if the following rank conditions are satisfied: $H_r : \Pi = \alpha\beta'$ of rank 0<r<p, where α and β are matrices of dimension p×r. β is a matrix representing the cointegrating vectors which are commonly interpreted as long-run equilibrium relations between the Z_t variables, while α gives the weights of the cointegration relationships in the VECM equations. The cointegration rank is usually tested by using the maximum eigenvalue (λ-max) and the trace test statistics proposed by Johansen (1988).

Identification and hypothesis testing on the cointegration space

The estimation of a VECM (1) subject to rank restrictions on the long-run matrix Π does not generally lead to a unique determination of cointegrating relationships and is not meaningful from an economic point of view. In fact, the structural estimation of the two cointegrating vectors obtained in the previous section requires the imposition of at least r^2 restrictions[7]. In recent years, considerable attention has been paid to the problem of identifying the long-run relationships in a linear cointegrating model. Johansen and Juselius (1994), Johansen (1995a), and Boswijk (1995), among others, have developed a testing procedure to identify cointegrating vectors by imposing linear restrictions in order to determine long-run behavioural parameters such as supply and demand elasticities.

However, sometimes it is more interesting to test joint restrictions on both the cointegration vectors and the adjustment coefficients. Johansen and Juselius (1990, 1992) developed a procedure to carry out individual tests on parameters from both matrices[8]. Mosconi (1998), extended the previous procedure to jointly consider general linear restrictions on both the long-run parameters, α and β. A general formulation of the null hypothesis can be expressed as:

$$\begin{aligned} \mathrm{H}0: \beta &= [\beta_1 \ldots \beta_r] = [H_1\,\varphi_1 \ldots H_r \varphi_r] \\ \alpha &= [\alpha_1 \ldots \alpha_r] = [\mathrm{A}_1 a_1 \ldots \mathrm{A}_r a_r] \end{aligned} \tag{2}$$

where: H_j is a ($k \times s_j$) matrix defining linear restrictions that reduce the k-dimensional vector β_j to the s_j-dimensional vector φ_j, with s_j representing the number of unrestricted parameters in β_j; k_j is the number of restricted parameters in β_j, such that $k_j + s_j = k$; similarly, A_i are ($k \times f_i$) restriction matrices α_i' s, where f_i is the number of unrestricted parameters in α_i.

Note that in the case where α is not restricted (A_i= I), (2) can be used to test the identification restrictions on β. In this case, the hypothesis is formulated as $\beta = (H_1\varphi_1, \ldots, H_r\varphi_r)$. As shown in Johansen (1995b), inference on the coefficients of cointegrated VAR systems is asymptotically based on mixed Gaussian distributions, so the Likelihood Ratio (LR) statistic for testing the hypothesis (2) is asymptotically $\chi^2(v)$. The degrees of freedom (v) can be calculated using the following expression:

$$v = \sum_{i}^{r} [k_i - (r-1)] + \sum_{i}^{r} (k - f_i)$$

Short-run dynamics: impulse response functions

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 should be calculated from the Moving Average Representation of the VECM (see Lütkepohl, 1993 and Pesaran and Shin, 1998):

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

where matrices B_i (i=2,…,n) are recursively calculated using the following expression: $B_n = \Phi_1 B_{n-1} + \Phi_2 B_{n-2} + \ldots + \Phi_k B_{n-p}$;B_0=I_p;B_n=0

for n<0; $\Phi_1 = I + \Pi + \Psi_1$; and $\Phi_i = \Psi_i - \Psi_{i-1}$ (i=2,…,k).

Following Pesaran and Shin (1998) the scaled Generalized Impulse Response Functions (GIRF) of variable Z_i with respect to a standard error shock in the j^{th} equation can be defined as:

$$GIRF\ (Z_{it}, Z_{j_t}, h) = \frac{e_i' B_h \Sigma e_j}{\sqrt{\sigma_{jj}}};\ \ h = 0, \ldots, n$$

where e_m (m=i, j) is the m^{th} column of the identity matrix (I_p).

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.

Results

The procedure outlined above has been applied to the system including the four variables described above (RER, RM, RAP, RAX). System (1) has been initially estimated including two lags with a constant term restricted in the cointegration space, implying that some equilibrium means are different from zero[9]. Moreover, in the case of Tunisia, in 1986 a Structural Adjustment Program was implemented which substantially changed the objectives and instruments of both the economic and agricultural policies. To account for this event on the level of the variables, model (1) was estimated including a restricted step dummy variable (D_t), which takes the unit value after 1986 and zero, otherwise. The multivariate autocorrelation (Godfrey, 1988) and normality (Doornik and Hansen, 1994) tests indicated that the model including the dummy variable was correctly specified[10].

Table 1 shows the results of Johansen's likelihood ratio tests for cointegration rank. At the 1% of significance level, both tests indicate that the null hypothesis of r=1 cannot be rejected. However, at the 5% of significance level, trace statistic indicate the existence of two cointegrating vectors. Note that since some dummy variables have been introduced, results of λ-max have to be interpreted with some caution (Johansen and al., 2001)[11]. For this reason, we have analysed the roots of the characteristic polynomial of the VAR model (called the companion matrix) since these provide useful information on how many (p-r) I(1) components exist in the data. As can be observed from Table 1 (the lower part), two values of the characteristic roots are close to unity, which confirms the presence of two cointegrating vectors in the system.

Table 1. Tests of the cointegration rank

Null hypothesis	*Trace*	*CV (1%)*[a]	*CV (5%)*[a]	*λ-max*	*CV (1%)*[b]
R = 0	76.16	70.41	63.42	32.14	33.24
R ≤ 1	44.02	48.25	42.50	20.70	26.81
R ≤ 2	23.31	29.63	25.24	16.21	20.20
R ≤ 3	7.10	13.82	11.23	7.10	12.97
Eigenvalues of the companion matrix	0.979	0.979	0.615	0.607	0.425

a. The critical values are taken from Johansen, Mosconi. and Nielsen(2001).
b. The critical values are taken from Osrterwald-Lenum (1992)

Identification of the long-run relationships

Taking into account the variables in the model, the following hypothetical cointegration relationships are expected:

$$RM_t - RAP_t + \beta_d^1 D_{1t} + \beta_c^1 = \mu_{1t} \tag{3}$$

$$RAX_t + \beta_{ER}^2 RER_t + \beta_d^2 D_{1t} + \beta_c^2 = \mu_{2t} \tag{4}$$

The first cointegration vector relates money supply and agricultural prices with the long-run homogeneity imposed. The second cointegrating relation represents a demand function for agricultural exports.

As a first step in the identification process, only restrictions on β parameters formulated in (3) and (4) are tested. The Likelihood Ratio (LR) statistic for testing the 4 over-identifying restrictions is 9.08, which is below the 1% critical value of $\chi^2(3)$ =11.34. Therefore, the null hypothesis cannot be rejected and the imposed restrictions have empirical support. The estimated restricted cointegration matrix is given by (estimated standard errors are given in parenthesis):

$$\beta' = \begin{pmatrix} -1.00 & 0.00 & 0.00 & 1.00 & \underset{(0.214)}{-0.71} & \underset{(0.701)}{1.46} \\ 0.00 & \underset{(0.311)}{-1.18} & 1.00 & 0.00 & \underset{(0.208)}{1.21} & \underset{(0.141)}{-0.55} \end{pmatrix} \times \begin{pmatrix} RM_1 \\ RER \\ RAX \\ RAP \\ 1 \\ D_1 \end{pmatrix} \quad (5)$$

Under these restrictions, the estimated parameters of the α matrix are:

$\hat{\alpha}$ matrix t-values for $\hat{\alpha}$

ΔRM	-0.055	0.058	1.523	2.833	
ΔRER	-0.110	0.051	-1.484	0.404	(6)
ΔRAX	-0.358	-0.371	-3.641	-3.405	
ΔRAP	-0.035	0.021	-2.621	1.116	

The first cointegration vector, interpreted as a long-run relation, indicates that a permanent increase in money supply leads to increased agricultural prices of the same magnitude. The second vector indicates that an increase in the exchange rate (a devaluation of the Tunisian currency) induces an increase in the agricultural exports, which is consistent with the economic theory. However, following Johansen (1995b) the dynamic properties of cointegrated VAR systems can be better understood by analysing the Moving Average Representation (MAR) of the process which will be considered in the next section.

Additionally, in this type of analysis, it is convenient to consider the estimated α_{ij} (i denotes the row and j the column) parameters, since they provide valuable information about the speed of adjustment of each variable towards the long-run equilibrium. In this paper we have checked which variables can be considered as the stochastic trends that are driving the dynamics of the system. From (6) it can be observed that the exchange rate does not react to the second long-run relationship and, thus, we have tested if it can be considered weakly exogenous in the long-run. Following Mosconi (1998), this hypothesis has been tested, jointly with the imposed restrictions on β (described in 5), by using the general hypothesis framework described in expression (2). The test value is 22.18, which is significant at the 1% level of significance (critical value, $\chi^2(6)$ = 18.5). Therefore, we reject the null hypothesis that the exchange rate is weakly exogenous.

Sensitivity of impulse response functions to alternative VAR specifications

As mentioned in the introduction, the main objective of this paper is to investigate the sensitivity of impulse response function analysis to alternative specification of VAR models. In other words, what are the consequences, in terms of policy analysis, of ignoring both the stochastic properties of the data and the long-run information? In order to test this, the IRF are calculated using the following three models: i) a VAR(2) model in levels; ii) a VECM without imposing any restriction on the cointegration vectors (UVECM), and iii) the VECM estimated in the last section with restrictions imposed on the β and α matrices (RVECM). Therefore, considering the GIRF for all the models, the idea is to measure the consequences of using what may be an incorrect model specification.

To facilitate the comparison between the different alternatives already mentioned, we have considered only the impulse response functions of agricultural variables to shocks on the macroeconomic variables. For all GIRF, the standard deviations are computed following Pesaran and Shin (1998)[12]. Figure 2 shows the response of agricultural variables to a one standard deviation shock to the exchange rate and money supply (the estimated two standard error boundaries are depicted as dashed lines).

Let us first consider a shock to the exchange rate. The response of agricultural exports in the three models has the expected positive sign, since a positive shock in the exchange rate means national currency devaluation and exports increase. However, the response in the VAR model the response is transitory (starting in the second period) while it is permanent in the other two models. Finally, comparing the UVECM and the RVECM, in the short run (first year) a shock in the exchange rate generates similar effects in agricultural exports. In both models there is an immediate response of the same magnitude which decreases slower in the RVECM.

In relation to price, the responses are quite different. The short-run response in all models is positive as expected. For the VAR model the price response is insignificant. For the UVECM, the exchange rate devaluation causes an increase in agricultural prices during two periods, returning to zero after that. In the case of RVECM, the very short-run response of agricultural prices to a positive shock in the exchange rate is of similar magnitude than in the UVECM but the effect is permanent. After 4 periods (years) agricultural prices reach their long-run state, remaining constant and significant. This last model reflects more accurately the current situation in Tunisia as, after 1986, the Government has tried to increase agricultural prices according to expected inflation so as to keep consumer prices constant in real terms.

A shock to money supply also generates different responses depending on the model specification. Again, the model in levels exhibits greater differences while for the other two models differences are limited to the very short run. In the case of the RVECM, as long-run money neutrality was imposed, price responses are consistent with such pattern. After four periods the long-run equilibrium is achieved. In the other two cases, the initial effect is not significant but become positive and significant after three periods. In the UVECM, also money neutrality is achieved although one period earlier than in the RVECM[13]. In the VAR model responses exhibit an upward trend indicating that long-run money neutrality does not hold. As can be observed, clearly policy implications should be rather different.

The response of agricultural exports to a money supply shock is quite similar across all models. The impact is quite small and non-significant in all cases, indicating that the main source of fluctuation in agricultural exports is exchange rate changes.

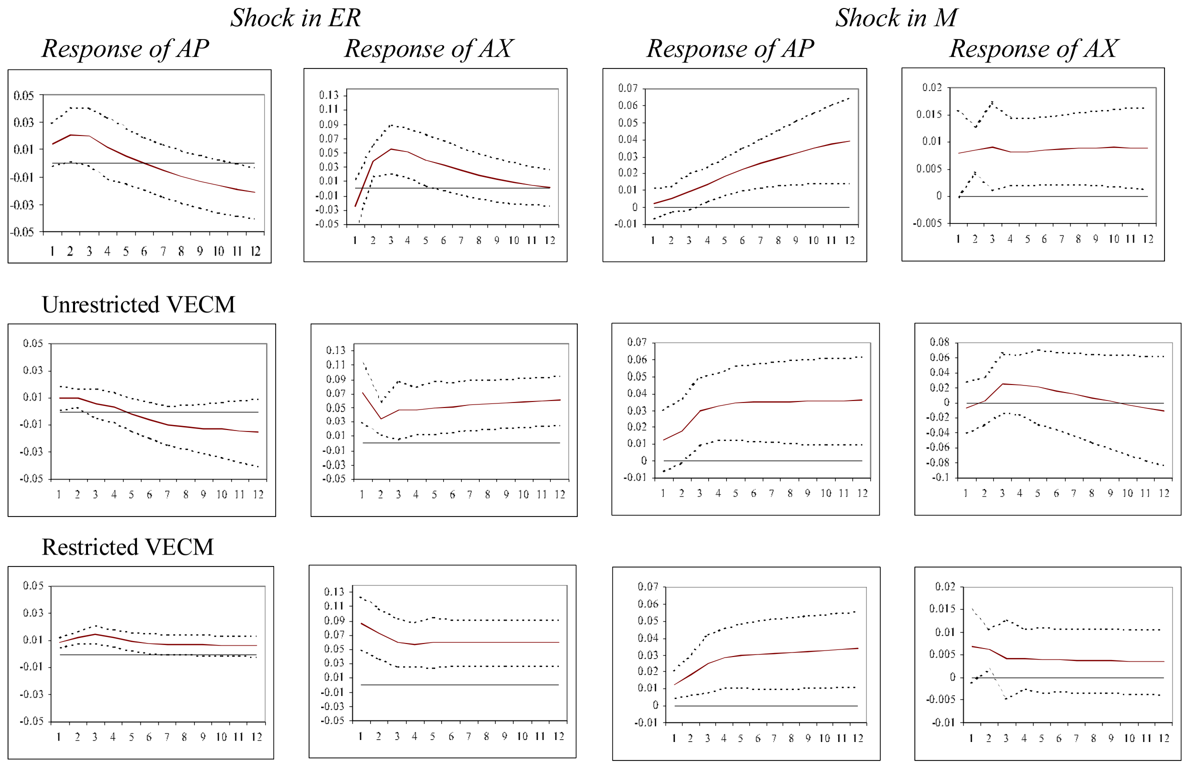

Figure 2. Generalised Impulse Response Functions VAR in level

Concluding remarks

The objective of the paper has been to analyse some of the most important relationships among macroeconomic variables and the agricultural sector in Tunisia. From the methodological point of view, the aim of this paper has been to provide an example of how alternative model specifications can lead to differing results, and then to different policy implications. When considering dynamic relationships between a set of economic variables, it is not only important to pay attention to the stochastic properties of the data (now a common practice) but also to the proper specification of the model. If variables are cointegrated, long-run relationships have to be correctly identified according to economic theory.

In the case of Tunisia, if cointegration is not considered, shocks on macroeconomic variables are assumed to have only a transitory effect on agricultural variables. Hence, any policy decision based on this assumption would generate unanticipated responses. In spite of recognising the limitations of this paper due to data availability we have tried to show that when analysing dynamics among a set of variables is important to properly identify the long run. Obviously, results reported her have to be examined in the light of both the chosen variables and the sample period used but they can be easily generalised to further empirical work.

Notes

1. See In and Mount (1994) for a literature review of this topic.
2. Stock and Watson (2001) provide a useful review of VAR modelling.
3. A multilateral rather than a bilateral real exchange rate would have provided better information about Tunisia's competitiveness in the trade market and on agricultural exports dynamics. However, these data are only available since 1983. In any case, when comparing the evolution of both rates since the information is available, it seems not to exist significant differences in relation to trends and turning points.
4. The objective of the monetary policy in Tunisia has been to keep inflation close to that of its main competitors. Traditionally, the government establishes the growth rate of the money supply 2% lower than the expected growth of the GDP. However, this objective has been always subject to revision within each year.
5. Agricultural exports in Tunisia mainly refer to olive oil (40% of total agricultural exports), fish (20%); dates (10%) and citrus (3%). Around 60% is exported to the European Union (EU). Except for the dates, the EU set a maximum amount to be imported with lower tariff. Only in the case of the olive oil, in some years total exports have exceeded the maximum amount allowed. Theoretically, the Tunisian government does not subsidised exports directly. However, from the 60's to mid 80's the agricultural policy was based on guarantee prices and subsidies for agricultural inputs for selected agricultural products (mainly food staples but also for some exporting products subject to export).
6. Results are not shown due to space limitations. They are available upon request.

7. When we impose (r-1) restrictions and one- normalization on each cointegration relation we say that β is just identified. In this case, no tests are involved because just-identifying restrictions do not change the likelihood function.
8. The general procedure is to test restrictions on the β parameters and afterwards on the α coefficients with the restrictions on β being imposed.
9. The lag length has been determined by both the Akaike and Schwarz Information criteria. With respect to the deterministic components, and following Harris (1995), several tests have been conducted to empirically select such components. Results indicated that a model with a restricted constant was statistically preferred.
10. The autocorrelation test was 11.54 with the critical value $\chi(16)$=28.85 while the multivariate normality test was 8.26 with the critical value $\chi(8)$=15.51 both at the 5% level of significance.
11. Several simulation studies have shown that the size distortion of the asymptotic LR tests for hypothesis testing on the cointegration vectors can be considerable in small samples (Abadir et al., 1999; Gredenhoff and Jacobson, 2001; and Johansen, 2000). Gredenhoff and Jacobson (2001) suggest the use of bootstrapping techniques, while Johansen (2000) suggests the use of Barlett corrections to reduce size distortions. A conservative alternative is to reduce the level of significance, which is similar to the traditional adjustment method found, for instance, in Reimers (1992). In this paper, since the sample period is relatively small, all tests are carried out using the 1% level of significance.
12. We have to note, however that in Pesaran and Shin (1998: pp. 27) expression (A.6) should be replaced by:

$$\begin{pmatrix} I_k & -I_k & 0 & \cdots & 0 \\ 0 & I_k & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & \ddots & 0 \\ 0 & \cdots & \ddots & I_k & -I_k \\ I_k & 0 & \cdots & \cdots & 0 \end{pmatrix}$$

13. This result indicates that the price parameter in the first cointegration relationship in the UVECM was close to one. In fact, it was 1.112. Results from the unrestricted model were not shown due to space limitations but are available from authors upon request.

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