Supply of Money and Food Prices: A Time Series Analysis

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Abstract

Relationships between monetary variables and price indices continue always to be the subject of research interest and studies. This paper examines the relationship between money supply and retail food prices in Greece, using individual time series of monthly data for these variables. ADF unit root testing shows that both series are non stationary at their levels. However, the series are stationary at their first differences and further analysis shows that the two I(1) variables are cointegrated, having a stationary, proportional, long-run equilibrium relationship. Both, the Johansen and Engle-Granger procedures are implemented. Estimation of Vector Error Correction (VEC) models allows for the derivation of the cointegrating vector and relationship, and results seem to justify the argument of money neutrality with regards to food prices. VEC estimation makes feasible also, the calculation of the adjustment speed to the long-run equilibrium between the two variables considered.

Q 110, E 510, E 310

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Introduction

The existence and nature of relationships between money supply and price indices has been a quantitative research issue, in order to assess the impact of monetary policies and liquidity on individual price indices or the relative prices between different groups of commodities. This is also true for prices related to agricultural activity, such as farm producer prices, marketing costs, prices paid by farm producers for inputs, and consumer food prices.

Alternative theories have accompanied relevant findings. According to the "cost-price squeeze hypothesis", given the oligopolistic nature of the farm input industries, the inflationary results of an expansionary monetary policy would lead to deterioration of agriculture's terms of trade. This is because prices paid by producers rise faster than prices received (Tweeten and Griffin 1976, Tweeten 1980, 1989). Moss (1992) has also examined the cost-price squeeze hypothesis, using cointegration analysis. Other authors (Bordo 1980, Rausser, Chalfant, and Stamoulis 1985, Frankel 1986) argue that expansionary monetary policy favors agricultural relative prices while the opposite occurs when monetary policy is contractionary. The underlying assumption is that the farming sector operates under more competitive conditions and price flexibility, which may result in short run price overshooting above levels of long run equilibrium when money supply rises. However, money maintains its neutrality in the long run, once adjustments to changes in money stock have been completed.

Greater sensitivity of agricultural prices to monetary shocks has also been found in Choe and Koo (1993). The same study finds a long run equilibrium relationship between money supply, agricultural prices and manufacturing prices, using a three variable Vector Autoregressive Model (VAR). However, no relationship was established between money supply and each one of the other two variables separately. Some uses of VAR models or impulse response functions derived from them, found varying degrees of response of agricultural prices to monetary changes and non neutrality of money for agricultural prices in US (Chambers 1984, Orden 1986, Devadoss and Meyers, 1987), Canada (Taylor and Spriggs 1989), and Brazil (Bessler, 1984). These results, based on traditional time series techniques and ignoring the long run behavior of the examined variables, are questionable according to Robertson and

Orden, (1990) who found evidence of money neutrality with respect to agricultural prices in New Zealand.

Since Cairnes (1871), the issue of factors affecting the speed of adjustment of agricultural prices to a changing supply of money was dealt with in several studies (Bordo 1980, Han, Jansen, and Penson 1990, etc). In the mentioned studies of Choe and Koo (1993) and Chambers (1984), agricultural prices were more sensitive than manufacturing prices to monetary changes in the short run. This agrees with Starleaf, Meyers, and Womack (1985). In the mentioned study of Robertson and Orden (1990), monetary shocks favored agricultural prices relatively to manufacturing prices in the short run, and raise permanently nominal prices. Similar long run results were caused by manufacturing price shocks, which however, in the short run were causing cost-price squeeze conditions for agriculture. Agricultural price shocks did not have such a significant impact on other prices. Using an imperfect information, rational expectations model for US, Lapp (1990) finds no evidence of serious monetary effects on relative prices of agriculture. Denbaly and Torgenson (1993) used cointegration analysis to study the existence of a relationship, as well as long and short term effects, of some macroeconomic variables on relative prices (farm to non farm) for wheat producers.

In fewer cases, including some of the above (eg. Bessler 1984), monetary effects on retail food prices have been considered. In general such studies establish the significant relationship between money supply and food prices (Belongia and King, 1983, etc). In the case of Greece, the relationship between monetary, other macroeconomic variables, and farm producer prices has been considered in Loizou, Mattas and Pagoulatos (1997). Non cointegration was found between supply of money and producer prices and a long term relationship cannot be established between these variables, but cointegration was established when other macroeconomic variables where simultaneously included in the analysis. However, there were only annual data for 27 years available for this study. In Karfakis (2002, 2004) evidence is provided for the validity of the quantity theory of money with regards to nominal national income in Greece.

This study focuses exclusively on the relationship between supply of money and consumer food prices in Greece. Given the significance of food expenditures for the consumer budget, it is important to know if a long term relationship connecting these variables exists, what its form is, and if money neutrality can be established. Unit root testing procedures for the two variables are followed by cointegration analysis, estimation of the cointegration vector, and a Vector Error Correction model (VEC), to derive the speed of short and long run adjustments to the long run equilibrium.

Data and Variables

Most studies referenced above, utilized quarterly or even annual data. In this study monthly time series were used, covering the period beginning in January 1970 and ending in December 2000. From January 2001 Greece is another EU member that joined the euro-zone, and the concept of a national supply of money no longer exists in its case. However, the 372 observations of a 31-year period do provide us with substantial information on the relationship between liquidity and food prices in Greece, which adds also to results of similar studies elsewhere.

Statistical data on money supply and retail food prices were provided by the Bank of Greece and the National Statistical Service. Monthly data on money supply for the whole period examined, were available for the value of circulating bank notes, coins, and demand deposits (M1), while information on consumer food prices was given by the Food Price Index (FPI).

Both series present similar behavior, that is, a general similar upward trend which is quite smooth until early 80's, but later the trend becomes and remains sharper despite some interim fluctuations. Towards the end of the period examined, the upward trend presents some signs of alleviation for both series. Smoothing the series by taking their logarithms, makes again clear the signs of non-stationarity. The series are negatively asymmetric (skewness) with positive kurtosis and the Jarque-Berra test rejects the null hypothesis at all levels of significance.

Unit Root Tests and Johansen Cointegration Analysis

Both variables and their monthly time series are used in their logarithmic form (LM1, LFPI). The Augmented Dickey-Fuller testing procedure (ADF) with a trend variable was implemented to test for unit roots in the series. The two AR(n) models used for this purpose take initially the form:

$$\Delta Y_t = \delta + \beta Y_{t-1} + \gamma t + \sum_{i=1}^{n} \alpha_i \Delta Y_{t-i} + \varepsilon_t \qquad (1)$$

where Y is LM1 and LFPI in the two equations respectively, t is a trend variable, for the white noise errors $\varepsilon \sim$ iid $N(0, \sigma^2)$, and all else is the parameters of the two equations.

In repeated regression estimates with various lag structures, the Akaike Information Criterion (AIC) and the Schwartz Bayesian Criterion (SBC) were used to adopt the lag length of the model which was found to be twelve for both time series. Moreover, the Lagrange Multiplier (LM) test of Breusch-Goldfrey and the Ljung-Box Q-statistic test confirmed the absence of residual autocorrelation at this lag structure.

Using (1) for both series, the ADF testing procedure showed that the variables are Difference Stationary Processes (DSP) rather than Trend Stationary Processes (TSP) and the use of a trend variable was rejected. Hence, we subsequently estimated for the ADF test, the following models without a trend variable:

$$\Delta LM1 = \delta_1 + \beta_1 LM1_{t-1} + \sum_{i=1}^{12} \alpha_i \Delta LM1_{t-i} + u_t \qquad (2)$$

$$\Delta LFPI = \delta_2 + \beta_2 LFPI_{t-1} + \sum_{i=1}^{12} \alpha_i \Delta LFP_{t-i} + \nu_t \qquad (3)$$

where δ's are constants, the last terms are errors, and the rest are the coefficients. Finally, results of the ADF test application, based on Dickey and Pantula (1987) and Dickey, Hasna, and Fuller (1987) were derived and are provided in Tables (1a) and (1b). (We did follow also the ADF testing procedure suggested by Dolado, Jenkinson, and Sosvilla-Rivero (1990), which in addition to providing the same result, confirmed the inclusion of a drift (constant) term in (1) for both series).

Table (1a) shows the unit root testing results for the levels of the variables, while Table (1b) includes the corresponding results when the first differences of the two variables are used instead. The last two columns include the results of the two tests for residual autocorrelation.

Table 1a: Unit Root Tests: Levels of variables

Variables	Lags	ADF test	LM (11)	Qstat (36)
LM1	12	-1.657497	2.921985	30.725
LFPI	12	-1.987753	13.83978	43.527

Comparing the ADF results with the DF critical values (-2.87 at 5% and -3.45 at 1% levels of significance) shows that there are unit roots and both variables are non stationary.

Table 1b: Unit Root Tests: First differences of variables				
Variables	Lags	ADF test	LM (11)	Qstat (36)
LM1	11	-6.346122	3.193087	34.564
LFPI	10	-4.591444	13.56696	43.599

Results in Table (1b) show that the first differences of variables are not characterized by unit roots and are stationary. It is concluded therefore, that both time-series of the variables are I(1).

We consider the VAR model for the two variables, given by:

$$Y_t = \sum_{i=1}^{k} A_i Y_{t-i} + \gamma + e_t \qquad (4)$$

where again $Y_t = [LM1_t, LFPI_t]'$, and each A_i is a (2x2) matrix of coefficients, and γ is a (2x1) vector of constant terms. Then, expressing (4) as a VEC model we have:

$$\Delta Y_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-1} + \upsilon_t \qquad (5)$$

Each Γ_i and the equilibrium or impact matrix Π are (2x2) coefficient matrices such that $\Gamma_i = -\sum_{j=i+1}^{k} A_j$ and $\Pi = \sum_{i=1}^{k} A_i - I$, while υ_t is the (2x1) error term vector. In each of the two equations described by (5) errors satisfy the usual assumptions and are non autocorrelated but they can be correlated across equations.

If the rank r of Π is $r(\Pi) < n$ (where now n=2) the two I(1) variables are cointegrated with the number of linearly independent cointegrating vectors r determining the degree of cointegration. There can be at most n-1 linearly independent cointegrating relationships which means, that in our case this relationship is unique. Moreover, the VEC rather than the VAR model should be estimated in this case. If cointegration exists, a decomposition of Π such that $\Pi=\varphi\psi'$ is possible, where φ and ψ are now (2xk) matrices such that $\psi' Y_{t-1}$ is stationary providing us with the cointegrating relationship while the columns of ψ are the cointegrating vectors, (implying now, as mentioned, one only linearly independent cointegrating relationship). The cointegrating relationship represents the long run equilibrium relationship between the two variables and the elements of φ are the adjustment speed coefficients.

The information criteria AIC and SBC as well as the LR test, converged to the choice of a maximum lag length of 13 months. The testing process for cointegration proposed by Johansen (1988, 1991, 1995), was implemented in the VAR model for the two variables. The Trace Statistic version of the Johansen test was adopted (critical values are found in Osterwald and Lenum (1992)) and results are given below in Table (2).

Table 2. Johansen test results				
Eigenvalues	Likelihood Ratio	5% critical Value	1% critical Value	Assumptions Cointegrating Relationships
0.0722	30.4250	19.96	24.6	None
0.0010	3.5948	9.24	12.97	At most one

Lack of cointegration is rejected at both levels of significance and the two I(1) variables are cointegrated C(1,1) implying r(Π) =1 since $r(\Pi) < n$ and r(Π) = 0 is also impossible given the Johansen test results. Estimation of the VEC using LS to calculate the elements of φ and ψ is inappropriate. However, the Johansen (1991, 1995) process of reduced rank regressions and partial canonical correlation analysis, provides a ML estimation of ψ and therefore of cointegrating vectors and relationships as well. Here, with r = 1 as said, the column vectors of ψ yield a unique cointegrating vector and relationship. Normalizing, this long term equilibrium relationship between the two variables was estimated as:

$$LFPI = 1.16LM1 - 11.99 \qquad (6)$$

The standard error and t-statistics for LM1 are 0.384 and 3.020 respectively while the corresponding numbers for the constant are 10.407 and 1.152. The coefficient of LM1 is highly significant but the constant term fails to pass the significance test at the 5% level.

The estimated relationship and statistical parameters confirm that there is a positive long-run equilibrium relationship between the two variables while the closeness of the estimated coefficient of LM1 to one provides support to the argument of money neutrality. Since $\frac{dLFPI}{dLM1}$ equals also the elasticity of FPI with respect to M1, the results imply also that a 1% increase in M1 raises, in equilibrium position, the value of the food price index by 1.15%.

Engle-Granger Cointegration Analysis

In dealing with the existence and estimation of the long run equilibrium relations of just two variables, the alternative Engle-Granger approach can be easily implemented as well. Despite the general preference for the Johansen approach, estimation of and Engle-Granger VEC allows for the simpler, straightforward and simultaneous testing for cointegration, estimation of the cointegrating relationship and the speed of adjustment from deviation to the long-run equilibrium. In our case of the two equations, under cointegration, an Engle-Granger VEC can be expressed as:

$$\Delta LFPI_t = -p_1(LFPI_{t-1} - \lambda_1 LM1_{t-1} - \lambda_0) + \sum_{j=1}^{k}\theta_{1j}\Delta LFPI_{t-j} + \sum_{j=1}^{k}\zeta_{1j}\Delta LM1_{t-j} + \upsilon_{1t} \quad (7)$$

$$\Delta LM1_t = -p_2(LM1_{t-1} - c_1 LFPI_{t-1} - c_0) + \sum_{j=1}^{k}\theta_{2j}\Delta LFPI_{t-j} + \sum_{j=1}^{k}\zeta_{2j}\Delta LM1_{t-j} + \upsilon_{2t} \quad (8)$$

Where now k=12 was adopted using the same criteria as above, p_1 and p_2 are adjustment speeds, and the error terms in the brackets are the cointegrating relationship ($c_1 = \lambda_1^{-1}$) of the two variables.

Consistency of the system assumes cointegration between LFPI and LM1 which implies that the error terms inside the brackets are stationary, i.e. I(0). Hence, testing for stationarity of the bracketed terms, tests for cointegration as well. Since their true values are unknown we can test using their estimates, but the DF critical values for the ADF test are no longer valid and use of more appropriate ones is necessary. However, following Banerjee et al (1986) for the single equation Error Correction Model (ECM), we can estimate directly equations (7) and (8) and test for stationarity of their error terms instead.

The general form of (7) under estimation, without the negative signs from p's and the bracketed terms, becomes:

$$\Delta LFPI_t = p_1^*(LFPI_{t-1} + \lambda_1^* LM1_{t-1} + \lambda_0^*) + \sum_{j=1}^{12}\theta_{1j}\Delta LFPI_{t-j} + \sum_{j=1}^{12}\zeta_{1j}\Delta LM1_{t-j} + \upsilon_{1t} \quad (9)$$

where p_1^* is the adjustment speed and is now negative. The estimated value of p_1 is -0.0005 and its t-statistic 3.71, while the bracketed term is stationary and its coefficients estimates are: $\hat{\lambda}_1^* = -1.2184$ and $\hat{\lambda}_0^* = 13.3703$ with t-statistics -2.0190 and 0.83726 respectively.

From the estimated coefficients θ and ζ of (9) we are more interested in the latter since we are concerned with the relationship between M1 and FPI. Four of those coefficients are significant at the 5% and 1% levels, belonging to the four, five, six and nine month, lagged variables of LM1. They all have positive signs (even though in general there are, as expected, alterations in sign and one third of the lagged LM1 variables are negative). Four other coefficients of lagged LM1 variables become significant at higher than 5% levels and they are all significant at the 10% level of significance. An R^2=0.50 and an $\overline{R}^2 = 0.46$ were estimated as well. The Q-statistic rejected the hypothesis of residual autocorrelation at the selected lag length.

The long-term equilibrium relationship is derived from the bracketed term showing deviation from equilibrium and can be expressed based on our findings as:

$$LFPI_t = 1.22LM1_t - 13.37 \qquad (10)$$

This result in (10) is similar to the other cointegrating relationship (6) derived using the Johansen approach. The estimated coefficients of LM1 are close and significant in both relationships while the constant term is non-significant in both cases. The small absolute value of the estimate $\hat{p}_1$ shows that deviations from the long run equilibrium are being corrected by 0.0005 per month which reflects a relatively slow speed of adjustment and the use of frequent monthly data as well.

Conclusions

The relationship between the retail food price index and supply of money has been examined, based on a 31 year period monthly data for Greece. Unit root testing showed that the two time series are non-stationary but they are first difference stationary. Both, Johansen and Engle-Granger cointegration analysis showed that the two variables are cointegrated, having a stationary long-term equilibrium relationship. The cointegrating relationship was estimated and there is a significant impact of money supply on consumer food prices. Moreover, the cointegrating relationship - especially the one estimated using the Johansen approach – seems to provide some support for the money neutrality hypothesis with respect to food prices. Food price changes caused by monetary policy are proportional in the long run to changes in money supply. This result agrees with some of the referenced studies above but disagrees with others, in other country cases. The time lags required for the full impact of money supply changes on consumer food prices were estimated. The speed of adjustment from deviation to the long-run equilibrium position was also calculated. The result implies a gradual, slow adjustment, less speedy than usual estimates, reflecting also the fact that calculated adjustments and lags are monthly.

Our study covers a long period which necessarily ends with Greece's entrance to the eurozone. In addition, there are shortcomings in testing and estimation procedures using time-series and this includes for example, the widely used in the literature Q-statistic (even though the magnitude of the impacts on the reliability of results is questionable). Such a study can be further pursued and expanded with other recent developments in time series analysis. However, the length of the covered period and frequency of data, as well as the result of the completed Engle-Granger and Johansen approaches and the similarity of their results with regards to the existence and estimation of a long-run equilibrium, does provide us with useful information.

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