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Food Price Volatility and Macroeconomic Factors: Evidence from GARCH and GARCH-X Estimates

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This article examines food price volatility in Greece and how it is affected by short-run deviations between food prices and macroeconomic factors. The methodology follows the GARCH and GARCH-X models. The results show that there exists a positive effect between the deviations and food price volatility. The results are highly important for producers and consumers because higher volatility augments the uncertainty in the food markets. Once the participants receive a signal that the food market is volatile, this might lead them to ask for increased government intervention in the allocation of investment resources and this could reduce overall welfare.

Key Words: relative food prices, volatility, macroeconomic factors, GARCH and GARCH-X models

JEL Classifications: E60, Q10, Q19

Over the recent past, food prices increased dramatically, leading to much debate about "the end of cheap food" period. This type of food crisis has serious implications for ecological sustainability, for the role of international financial institutions, and for the risk of future nutritional emergencies. In rich countries, food covers a relatively small part of a household's budget; by contrast, in poor countries, households use a large share of their income for food expenses, implying that food price increases lead to reduced real income as well as to higher risks of malnutrition and higher uncertainty (volatility) in food markets, because food price inflation severely stresses the most vulnerable groups. Nevertheless, these international prices for the major food types have decreased almost just as dramatically as they had increased, exacerbating the magnitude of food price volatility. This decline in food prices accompanies the dramatic fall in international economic activity resulting from the global economic slowdown. However, a report by the FAO (2009) shows that food prices have remained "sticky" in many countries, implying that they remain at high levels. As a result, many low-income countries and, especially, households continue to be adversely affected by high levels of food prices.

Factors such as the role of financial speculation in food commodity markets along with global financial markets' turmoil, export bans, adverse weather conditions, precautionary demand for food stocks, lack of efficient logistics systems, infrastructure for food marketing and

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distribution, rising energy prices, and energy intensity of the agricultural sector, the diversion of certain food commodities to produce alternative fuel, and political factors through policy inadequacies, weak institutions that undermine incentives for agricultural production, input subsidies, and involvement of public agencies in food imports have received criticisms about their contribution to such an uncertain food market environment. Certain empirical studies identify that unexpected trading volumes in commodity futures trading lead to higher cash price volatility (Sahi and Raizada, 2006; Yang, Balyeat, and Leatham, 2005), whereas Gilbert (2008) finds weak evidence that such speculative activities could influence food prices. Campiche et al. (2007) and Thompson, Meyer, and Westhoff (2009) report evidence in favor of the impact of oil prices on production costs, whereas the reverse route is nonoperative. Furthermore, Irwin and Good (2009) provide evidence that a new era of crop price volatility has begun with considerable uncertainty about the new level of average nominal prices causing great stress to market participants. They also state that the change in crop markets today is comparable to those during the mid-1970s and they anticipate that market participants will adjust to the new pricing environment with surprising speed. Von Braun and Torero (2009) identify two major explanations behind the 2007–2008 international food price crisis; first, the ad hoc trade policy interventions such as export bands or high export tariffs or high import subsidies, and second, the significant flow of speculative capital from financial investors into agricultural commodity markets. They also suggest several changes in regulatory frameworks to reduce speculation in food commodities. Finally, Harri, Nalley, and Hudson (2009) report that not only higher costs of producing energy from food products, but also greater use of oil-based inputs in food markets have generated higher food prices. By contrast, Yu, Bessler, and Fuller (2006) provide empirical evidence that oil price shocks have a relative small effect on food prices. The presence of mixed results is an argument against using oil prices as an explicit determinant of food prices in this study.

Conceicao and Mendoza (2009) argue that lack of investments in the agricultural sector is the most critical factor for food price increases. Roache (2010) finds that the macroeconomic variables that really matter for food price volatility are inflation and exchange rates. In addition, he argues that the presence of such high food price volatility has made the required policy responses more challenging, whereas it seems to have added an extra nebulous to investment and consumption decisions made by businesses and households, respectively.

This high increase in food prices (both in mean and in volatility) is eliciting policy responses that exacerbate rather than cushion price volatility as policymakers rush to restrict exports, control domestic prices, and attempt to rebuild stocks in the face of price increases. Macroeconomic instability has real sectoral effects, but it retains the case that there is no predictable direction in which real farm prices are affected by general inflation. Grennes and Lapp (1986), using annual data, consider the extent to which macroeconomic factors that generate inflation systematically alter relative agricultural prices. They could not reject the hypothesis that real aggregated agricultural prices have not been altered by the level of money prices, inflation, or exchange rates. They argue, however, that the use of monthly data, for example, might show temporary relative price effects, as is shown by the studies of Chambers (1984), using monthly data, and Chambers and Just (1982), using quarterly data. Gardner (1981) and Grennes and Lapp (1986) argue that an economic environment that generates changes in its components is expected to trigger substantial price swings, whereas Chambers (1985) exemplifies the fact that food products are less storable than nonfood products, resulting in higher relative price variability. Blisard and Blaylock (1993) also report empirical results that lend support to the argument that food price volatility exhibits strong swings over time and, thus, these prices should not be included in the estimation of general price inflation, i.e., only "core" inflation matters for economic policy evaluation.

Recent contributions on food prices have also emphasized the role of certain macroeconomic factors and policies such as monetary, fiscal, trade, and exchange rate policies in the formulation of agricultural price policy (Gray, 1992; Kargbo, 2000). Barnett, Bessler, and Thompson (1983), Chambers and Just (1982), and Schuh (1986) demonstrate that the increases of food prices in the U.S. economy during the 1970s are considerably the result of macroeconomic factors. The identification of such factors increases the efficiency in the use of inputs of production as well as the availability of food products, allowing final consumers to purchase those food products at affordable prices.

Meyers et al. (1986) and Taylor and Springs (1989) exemplify the role of exchange rates in determining agricultural prices. Exchange rates can affect food prices mainly through the mechanism of international purchasing power and the effect on margins for producers with non-U.S. dollar costs (Gilbert, 1989). Taylor and Springs (1989) and Tegene (1990) find substantial effects from monetary factors on agricultural prices, whereas Kliesen and Poole (2000) argue that monetary policy can affect the agricultural sector only in an indirect way by contributing to low inflation, stable inflation expectations, and low interest rates. By contrast, Bessler (1984), Isaac and Rapach (1997), Orden (1986), and Orden and Fackler (1989) show that monetary impacts were not the dominant factors for food prices. Interest rates can also affect food prices, especially if market participants expect interest rate shocks to persist (Frankel, 2006).

Other empirical studies have already identified a relationship between expected inflation and changes in relative prices of particular products (Ball and Mankiw, 1992; Lach and Tsiddon, 1993; Loy and Weaver, 1998; Mizon, 1991; Smith and Lapp, 1993; Stockton, 1988). Moreover, higher food prices mean higher inflation and, given the large weight of food prices in the Consumer Price Index, inflation will rise as a result of persistent food price increases, leading to higher wages and causing inflationary expectations to become embedded in economies. Inflation is also expected to reduce real consumption, savings, and investments, all of which may combine to slow down aggregate demand and, thus, to dampen economic activity. Calvo (2008) and Frankel (2008) argue that the link between food prices and macroeconomics is mainly the result of the global financial crisis linked with excess liquidity in global economies and nourished by low interest rates, especially set by the G7 central banks along with high economic growth figures in China and India. Attie and Roache (2009) and Leibtag (2008) provide empirical support to the argument that commodities can be largely used as hedging investment portfolios against inflation risk. They exemplify the role of food price volatility in affecting the portfolio choices of financial investors.

The empirical analysis has concentrated on revealing any relationship between food prices and macroeconomic factors in terms of the means of the variables under investigation; however, in terms of volatility, it has remained rather silent. Macroeconomic factors could potentially affect food price volatility through certain channels, leading to persistent changes in supply and demand conditions in food markets. In case that these factors exemplify the volatility of food prices, this could create higher uncertainty about future food prices and, thus, participants in food markets and academics may attribute any changes in the mean and the volatility of food prices to such factors or events. This type of uncertainty about food prices, in turn, could affect both producers' and consumers' decisions as well as any investment activity in the food industry. In addition, if the volatility of macroeconomic factors does seem to play a substantial role in determining the volatility of food prices, then policymakers should take it into consideration in formulating price and income programs for the agricultural sector. With volatility at high levels, uncertainty delays or cancels investments and changes in consumption that would have been more likely with more stable prices.

Monetary factors, i.e., money supply, are linked to food prices through certain mechanisms such as the foreign exchange markets in which financial resources move among global economies. More specifically, the Greek economy is considered a small open economy well integrated on a global basis. As a result, monetary policy changes could affect food prices through real exchange rates that measure the external competitiveness of an economy. Thus, changes in such rates transform the structure of relative prices between tradable goods and nontradable goods with agricultural products belonging to the first group (Jaeger and Humphreys, 1988). The impact of fiscal policy activities on food prices is realized through an indirect mechanism. In particular, fiscal management affects not only the course of domestic interest rates (Gale and Orszag, 2003; Hauner and Kumar, 2006; Modigliani and Jappelli, 1988) as well as exchange rates (Canzoneri, Cumby, and Diba, 2001; Daniel, 1993, 2001; Dupor, 2000), but also domestic inflation and the course of current account; such fiscal actions could have a substantial impact on the demand for the products of various sectors in the economy, including agricultural products (Kargbo, 2005).

For the case of the Greek economy, certain degrees of inflation rates in the nonfarm economy, large shocks in inputs prices, e.g., feed grains, high competition from other European Union (EU) countries along with the changes of the Common Agricultural Policy (CAP), resulted in highly volatile food prices over the last two decades (Apergis and Rezitis, 2003). In particular, CAP reforms caused significant decreases in intervention prices and induced compensation to producers through direct payments, which are not related to the level of production, causing higher food price volatility. Moreover, GATT changes might have caused higher food price volatility, because they have fostered more intense international competition through limiting support programs to agriculture. The aforementioned arguments are supported by the study of Yang, Haigh, and Leatham (2001), which examines the effect of liberalization policy on agricultural commodity price volatility in the United States and finds that liberalization policies caused an increase in price volatility for wheat, soybeans, and corn.

Very few studies have investigated the role of macroeconomic factors in the process of relative food prices in Greece. Daouli and Demoussis (1989) find that over the period 1967–1987, price support policies resulted in neutralizing the impact of inflation on Greek agriculture. Our study, however, contributes to the relevant literature by extending such empirical work and investigating the impact of more macroeconomic factors on food prices. After all, inflation is always a phenomenon that reflects macroeconomic developments and such developments have been quite (positively) crucial in the Greek economy following the Maastricht Treaty. Varangis (1992), through the methodology of Autoregressive Conditional Heteroscedasticity (ARCH) models, finds that money supply developments strongly affect the volatility of food prices. Maravegias (1997) argues that the manner in which the exchange rate policy has been implemented, relative to the evolution of the general price index, seems to have had a direct effect on the prices Greek farmers received. At the same time, the efforts of Greek policymakers since the mid-1980s to reduce substantially high inflation—through the implementation of a strict monetary policy that led to high interest rates—have also exerted a substantial impact on food prices (Maravegias, 1997). Zanias (1998) identifies a role for macroeconomic factors that seems to have a substantial impact on Greek inflation, which, in turn, alters food prices. He claims that macroeconomic instability contributes to higher food price volatility, which in turn has adverse effects on agricultural production and income. Finally, Hondroyiannis and Papapetrou (1998) examine the relation between money supply and agricultural prices over the period 1972–1994. Their results show that money supply changes can affect agricultural prices through the mechanism of interest rates.

The objective of this article is to investigate the manner short-run deviations from the relationship between food prices and macroeconomic factors that drive food price volatility in Greece. Therefore, when volatility is increased as a result of shocks in the system, it is reasonable to investigate the behavior of conditional variance as a function of short-run deviations from the equilibrium path. The accurate measurement of food price volatility is important not only because volatility causes uncertainty to producers, consumers, and policymakers, but also costs are inevitably incurred (McMillan, 2003). A significant positive effect would imply that short-run deviations affect not only the conditional mean, but also the conditional variance, implying that the further food prices deviate from macroeconomic variables in the short run, the harder they are to predict. In other words, conditional heteroscedasticity could be modeled as a function of lagged error correction terms affected by macroeconomic conditions in case that disequilibrium measured by such error correction terms is responsible for uncertainty measured by the conditional variance. A significant effect indicates these terms have potential power in modeling the conditional variance of food prices. Therefore, last period's equilibrium error has a significant impact on the adjustment process of the relevant variables. In this case, a positive effect of the short-run deviations on the conditional variance implies that as the deviation between food prices and macroeconomic variables gets larger, the volatility of food prices increases and prediction becomes harder. Thus, given that short-run deviations from the long-run relationship between food prices and macroeconomic variables may affect the conditional variance, then they will also affect the accuracy of such predictions. The empirical findings would be of value to commodity market participants as well as to policymakers.

The methodology followed in this article to measure relative food price volatility is that of Generalized Autoregressive Conditional Heteroscedastic (GARCH) models introduced by Bollerslev (1986). Chou (1988) argues in favor of GARCH models on the grounds that they are capable of capturing various dynamic structures of conditional variance, of incorporating heteroscedasticity into the estimation procedure, and of allowing simultaneous estimation of several parameters under examination. Finally, the methodology of the GARCH-X models, introduced by Lee (1994), is also followed. This model allows the link between short-run deviations from a long-run cointegrated relationship and volatility. Whether it is stronger in providing better and more reliable forecasts is still an issue in dispute. Nevertheless, the literature offers a great variety of models that can be used to forecast food prices and food price volatility. According to Kargbo (2007), such techniques include exponential smoothing, ARIMA modelling, Vector Autoregression (VAR), and Vector Error-Correction Models. In general, forecasting food prices through an adequate model seems to be more than a necessity considering the importance of food policy reforms performed by many (especially, low-income) economies.

The remaining text of this article is organized as follows. The next section describes the methodology used, whereas the following section presents the empirical analysis and discusses the empirical results. The final section concludes the article.

The Methodology of GARCH and GARCH-X Models

The ARCH methodology, pioneered by Engle (1982), suggests a method for measuring uncertainty if it is serially correlated. The empirical methodology used here extends the ARCH model. Let ξ_t be a model's prediction error, b a vector of parameters, and x_t a vector of predetermined explanatory variables in the equation for the conditional mean:

$$(1) \qquad y_t = x_t b + \xi_t \quad \xi_t|\Omega_{t-1} \sim N(0,h_t)$$

where h_t is the variance of ξ_t, given information Ω at time $t - 1$. The GARCH specification, as developed by Bollerslev (1986), defines h_t as:

$$(2) \qquad h_t = u_0 + \sum_{i=1}^{p} c_i\, h_{t-i} + \sum_{j=1}^{q} a_j\, \xi^2_{t-i}$$

with u_0, a_j, and c_i being nonnegative parameters. The parameter u_0 indicates volatility acting as a floor, which prevents the variance from dropping below this level (Staikouras, 2006). According to equation (2), the conditional variance h_t is specified as a linear function of its own lagged p conditional variances and the lagged q squared residuals. Engle and Bollerslev (1986) and Lamoureux and Lastrapes (1990) have argued that if $\sum c_i + \sum a_j = 1$, then the GARCH specification turns into an integrated GARCH (IGARCH) process, implying that current shocks persist indefinitely in conditioning the future variance. Maximum likelihood techniques are used to estimate the parameters of the GARCH model according to the BHHH algorithm (Berndt et al., 1974). According to Lamoureux and Lastrapes (1990), the GARCH model may provide biased estimates of persistence in variance in case that additional information arising from other factors is unaccounted for. Therefore, we proceed to present an augmented version of the traditional GARCH model, which is known as the GARCH-X model. This model takes into consideration the effects of the short-run deviations on the conditional variance. In this respect, the specification equation (2) becomes:

$$(3) \qquad h_t = u_0 + \sum_{i=1}^{p} c_i h_{t-i} + \sum_{j=1}^{q} a_j \xi^2_{t-i} + \gamma_1 z^2_{t-1}$$

Moreover, in terms of persistence, like in the case of the GARCH model, for the GARCH-X model to be stationary, we need: $\sum_i c_i + \sum_j a_j < 1$. We also need: $u_0 > 0$ and all c_i and $a_j \geq 0$ for $i = 1, \ldots, p$ and $j = 1, 2, \ldots, q$. i...,j

The short-run deviations are denoted by the squared and lagged error-correction term $z_{t-1}{}^2$. This term, known as the Error Correction (EC) term, acts as a proxy for the residuals from a cointegrating vector that associates food prices and certain macroeconomic variables and it is believed to have important predictive powers for the conditional variance of food prices. The parameter γ_1 indicates the effects of the short-run deviations between the macroeconomic variables from a long-run cointegrated relationship on the conditional variance of the residuals of the food prices equation. If γ_1 is positive, this implies that food prices become more volatile and possibly less predictable as the deviation of food prices from certain macroeconomic factors gets larger. In other words, the presence of such deviations in the conditional variance may be exploited to get more precise and reliable confidence intervals for forecasts of food prices. According to Hwang and Satchell (2001), the GARCH-X model is simple but includes additional information, vis-à-vis the GARCH model, on some important factors such as the macroeconomic deviations of the economy as they are captured by the deviations from the macroeconomic equilibrium path.

Empirical Analysis

Data

The empirical analysis is carried out using monthly data on food prices proxied by producer prices (prices received by farmers), money supply defined as M1, income per capita (YPOP) defined as the ratio of industrial production index divided by the population index,[1] the real exchange rate (RE) measured as the real effective exchange rate index (1995 = 100), and the budget deficit (or surplus) to income ratio (DEFY). The effective exchange rate is defined in such a manner as a decline in the ratio to be consistent with a real appreciation of the domestic currency. Deficit, money supply, and food prices are divided by the Gross Domestic Product deflator. In this manner, the real deficit to income ratio (RDEFY), real money balances (RM), and relative food prices (PP), respectively, are obtained. Data span the period 1985–2007 and are obtained (except that on food prices) from the Research Department of the Bank of Greece. Food prices are described as an index of producer prices of agricultural products (1995 = 100) and they are obtained from the Eurostat NewCronos database. Note that an index of producer prices received by farmers for food items is not available for the period under consideration, whereas a price index was necessary to capture the price behavior of the whole food sector. Moreover, a price index at the producer price level was used and at the consumer price level because the former does not include imports.

The period coincides with the period in which Greece has been a full member of the EU (European Economic Community), whereas important economic policy incidents occurred. For empirical purposes of the study, a dummy variable is considered that is related to the reforms adopted with respect to the implementation of the CAP that occurred in May 1992. The reform of the CAP decreased price support for farmers and increased direct income support. As Ray et al. (1998) argue, however, our data set cannot be used to determine the portion of price variability that could be attributed to the CAP reform and policy changes instituted in the 1992 CAP Treaty.

The original data are seasonally unadjusted. However, a certain number of researchers have

[1] The per-capita GDP or GNP for Greece is available on an annual or quarterly basis. In the present article, we used monthly data to estimate the proposed models. To this end, we used the industrial production index instead of the per capita GDP or GNP, which is available on a monthly basis. The share of industrial sector in GDP of Greece is approximately 20.3%.

reported that using seasonally unadjusted data and then applying certain statistical techniques to account for seasonality, i.e., seasonal unit roots or seasonal dummies, generates incorrect signs or statistically insignificant estimates (Lee and Siklos, 1991; Osborn, 1990). Thus, our data set is seasonally adjusted through the X11 procedure. This procedure is based on the assumption that the original series are composed of seasonal, trend, cycle, trading date, and irregular components. The procedure estimates each component of the original series in an iterative process, which makes extensive use of moving averages and a methodology for identifying and replacing extreme values in the data set before providing final estimates of the components of the adjusted series. Throughout the article, lower case letters indicate variables expressed in logarithms. Finally, RATS6.1 software (provided by Estima, US) assists the empirical analysis.

Integration Analysis

We first test for unit root nonstationarity by using the Augmented Dickey-Fuller (ADF) test proposed by Dickey and Fuller (1981). The lag length is determined through the general-to-specific method presented by Perron (1997). Table 1 reports the results. The hypothesis of a unit root is not rejected for the variables of real deficit to income ratio, real money supply, the real exchange rate, income per capita, and food prices at the 5% significance level. When first differences are used, unit root nonstationarity is rejected for all variables under study.

However, the ADF test has received very strong unfavorable critique as a result of its low power, especially in small samples. Thus, to increase the power of our unit, root results alternative tests are also used such as the Kwiatkowski et al. (KPSS, 1992) test in which stationarity is the null rather than the alternative hypothesis. These results are also reported in Table 1. The KPSS results, using zero, one, two, three, and four lags, show that the null hypothesis of level stationarity and trend stationarity is rejected for all variables under study.

Moreover, to further check the robustness of unit root test, the efficient unit root tests, proposed by Elliott (1999) and Elliott, Rothenberg, and Stock (1996), are also reported in Table 1. These latter tests avoid the problem of short-spanned data. The lag lengths in both efficient tests remain the same as in the ADF test, whereas both versions are reported with and without trend. In all cases and for all variables under investigation, the empirical findings indicate that the null hypothesis of stationarity is rejected at the 5% significance level. Overall, there is consistency in our unit root testing and the presence of cointegration is valid to be tested.

Cointegration and Error Correction Analysis: A Mean Equation for Relative Food Prices—Cointegration Analysis: With and without Breaks

Once having identified a set of five jointly dependent stochastic variables integrated of the same order, i.e., I(1), a VAR model is postulated to obtain a long-run relationship. Tests developed by Johansen and Juselius (1990) revealed evidence in favor of cointegration. The results are reported in Table 2.

Both the eigenvalue test statistic and the trace test statistic indicate that there exists a single long-run relationship between relative food prices and the macroeconomic variables under consideration. The description of the cointegration space yields:

$$\underset{}{pp} = \underset{(4.47)^*}{0.369\ \mathrm{RDEFY}} - \underset{(-3.48)^*}{0.174\,\mathrm{rm}} + \underset{(3.72)^*}{0.146\,\mathrm{re}} + \underset{(3.96)^*}{0.105\,\mathrm{ypop}} + \underset{(3.18)^*}{0.0874}$$

$R^2 = 0.73$ LM = 10.44[0.27] RESET = 13.09[0.32], where pp denotes food prices, RDEFY denotes the public deficit to income ratio, rm denotes real money balances, re denotes the real exchange rate, and ypop denotes income per capita. LM is a serial correlation test

Table 1. Unit Root Tests

ADF Test	Without Trend		With Trend	
Variable	Levels	First Differences	Levels	First Differences
RDEFY	−1.83 (5)	−8.30 (4)[a]	−1.14 (5)	−8.42 (4)[a]
rm	−2.44 (7)	−5.25 (4)[a]	−0.73 (6)	−5.76 (5)[a]
re	−2.03 (4)	−6.38 (3)[a]	−2.16 (7)	−6.44 (4)[a]
ypop	−2.17 (6)	−10.51 (3)[a]	−1.94 (6)	−8.87 (4)[a]
pp	−2.42 (3)	−10.28 (2)[a]	−2.39 (4)	−7.21 (3)[a]

KPSS Test Variable	Level Stationarity	Trend Stationarity
RDEFY	1.1983 (l = 0)	1.3471 (l = 0)
	1.1864 (l = 1)	1.3277 (l = 1)
	1.1481 (l = 2)	1.2352 (l = 2)
	1.1184 (l = 3)	1.1684 (l = 3)
	1.1095 (l = 4)	1.1172 (l = 4)
Rm	1.1766 (l = 0)	1.2527 (l = 0)
	1.1468 (l = 1)	1.2239 (l = 1)
	1.1235 (l = 2)	1.1540 (l = 2)
	1.0842 (l = 3)	1.1236 (l = 3)
	1.0655 (l = 4)	1.1093 (l = 4)
re	1.2347 (l = 0)	1.3544 (l = 0)
	1.2094 (l = 1)	1.3093 (l = 1)
	1.1541 (l = 2)	1.2422 (l = 2)
	1.1288 (l = 3)	1.1981 (l = 3)
	1.1036 (l = 4)	1.1535 (l = 4)
ypop	1.1847 (l = 0)	1.2236 (l = 0)
	1.1437 (l = 1)	1.1874 (l = 1)
	1.1153 (l = 2)	1.1346 (l = 2)
	1.0964 (l = 3)	1.1058 (l = 3)
	1.0655 (l = 4)	1.0861 (l = 4)
pp	1.2095 (l = 0)	1.2879 (l = 0)
	1.1764 (l = 1)	1.2252 (l = 1)
	1.1533 (l = 2)	1.1879 (l = 2)
	1.1239 (l = 3)	1.1456 (l = 3)
	1.1052 (l = 4)	1.1233 (l = 4)

Elliott et al. Test	Without Trend		With Trend	
Variable	Levels	First Differences	Levels	First Differences
RDEFY	−0.79 (5)	−3.79 (4)[a]	−1.25 (5)	−4.15 (4)[a]
rm	−0.64 (7)	−3.11 (4)[a]	−0.96 (6)	−3.69 (5)[a]
re	−1.16 (4)	−4.57 (3)[a]	−1.47 (7)	−4.77 (4)[a]
ypop	−1.22 (6)	−4.66 (3)[a]	−1.71 (6)	−4.80 (4)[a]
pp	−1.38 (3)	−3.93 (2)[a]	−1.82 (4)	−4.38 (3)[a]

Elliott Test	Without Trend		With Trend	
Variable	Levels	First Differences	Levels	First Differences
RDEFY	−1.13 (5)	−4.33 (4)[a]	−1.52 (5)	−4.72 (4)[a]
rm	−0.86 (7)	−3.58 (4)[a]	−1.07 (6)	−4.13 (5)[a]
re	−1.41 (4)	−4.83 (3)[a]	−1.36 (7)	−5.66 (4)[a]
ypop	−1.68 (6)	−4.92 (3)[a]	−1.95 (6)	−5.38 (4)[a]
pp	−1.82 (3)	−4.29 (2)[a]	−1.99 (4)	−4.68 (3)[a]

[a] Significant at 5%.

Table 2. Cointegration Tests

r	n-r	m.λ.	95%	Tr	95%
Johansen-Juselius Test					
Lags = 10					
r = 0	r = 1	42.8940	37.0700	96.2427	82.2300
r ≤ 1	r = 2	29.2280	31.0000	53.3487	58.9300
r ≤ 2	r = 3	12.5041	24.3500	24.1207	39.3300
r ≤ 3	r = 4	8.5881	18.3300	11.6166	23.8300
r ≤ 4	r = 5	3.0285	11.5400	3.0285	11.5400
Saikkonen and Lütkepohl Test					
Lags = 9					
r = 0	32.47 [0.026]				
r = 1	4.58 [0.197]				
r = 2	1.39 [0.304]				
r = 3	0.95 [0.426]				
r = 4	0.27 [0.587]				

r, number of cointegrating vectors; n-r, number of common trends; m.λ., maximum eigenvalue statistic; Tr, Trace statistic. Figures in brackets denote *p* values. The number of lags was determined through Likelihood Ratio tests developed by Sims (1980).

and RESET is a model specification test. Numbers in parentheses denote *p* values. Both tests identify the adequacy of our model.

However, in May 1992, critical changes occurred in the CAP implementation. Thus, a dummy variable (with values of zero up to 1992:4 and one thereafter) that captures the restructuring conditions of this CAP policy change is also considered in the cointegrating vector. The Saikkonen and Lütkepohl (2000) test with a break at 1992:4 is used to test for any possible break shift of the cointegration vector. This test rejects the null hypothesis of no cointegration. The results of this test incorporating this shift are also reported in Table 2. These findings also provide evidence in favor of cointegration between the variables under study. The new cointegration vector yields:

$$\underset{}{pp} = \underset{(4.33)^*}{0.395\,\mathrm{RDEFY}} - \underset{(-4.05)^*}{0.213\,\mathrm{rm}} + \underset{(3.95)^*}{0.174\,\mathrm{re}} + \underset{(4.11)^*}{0.159\,\mathrm{ypop}} + \underset{(3.63)^*}{0.104}$$

$R^2 = 0.81$ LM = 12.56[0.31] RESET = 14.55[0.36]

To summarize, the analysis of the long-run structure of the data, in terms of stationary cointegration relations, provides the following findings:

1. The hypothesis that relative food prices are cointegrated with a set of macroeconomic variables such as real public deficit, real money supply, the real exchange rate and per capita income is accepted.
2. All macroeconomic variables exert a statistically and significant effect on relative food prices.
3. Real public deficits, the real exchange rate, and the per-capita income exert a positive effect on relative food prices, whereas real money supply exerts a negative effect on those prices.
4. Finally, these results are valid with and without the presence of a structural break occurred in May 1992 under the CAP restructuring.

The Short-run Structure

Having established the presence of a cointegrating relationship between relative food prices, on the one hand, and the real deficit-to-income ratio, real money balances, the real exchange rate, and income per capita on the other hand (and based on the equation that considers the structural break event), a parsimonious error correction vector autoregressive mechanism is considered, which adds the residuals from the cointegrating vector. The analysis yields the following estimates:

$$
\begin{aligned}
\Delta pp = {} & \underset{(5.54)^{*}}{0.389\,\Delta pp(-4)} + \underset{(2.77)^{*}}{0.12\,\Delta pp(-5)} + \underset{(2.93)^{*}}{0.212\,\Delta pp(-8)} + \underset{(4.61)^{*}}{0.494\,\Delta RDEFY(-1)} - \underset{(-3.07)^{*}}{0.0392\,\Delta RDEFY(-2)} \\
& + \underset{(2.17)^{*}}{0.0231\Delta RDEFY(-3)} - \underset{(-3.57)^{*}}{0.151\,\Delta rm(-4)} + \underset{(3.82)^{*}}{0.168\,\Delta re(-8)} + \underset{(2.15)^{*}}{0.078\Delta ypop(-3)} + \underset{(3.01)^{*}}{0.072\,Dypop(-4)} \\
& + \underset{(6.29)^{*}}{0.236\,\Delta ypop(-6)} + \underset{(3.94)^{*}}{0.154\,\Delta ypop(-7)} + \underset{(3.47)^{*}}{0.094\,\Delta ypop(-8)} - \underset{(-5.02)^{*}}{0.418\ EC(-1)} - \underset{(-2.86)^{*}}{0.19\ CAP}
\end{aligned}
$$

$$R^2 = 0.89 \text{ LM} = 1.53[0.22] \text{ RESET} = 0.019 \times [0.89] \text{ NO} = 3.14[0.21] \text{ HE} = 2.3[0.13]$$

$$\text{ARCH}(1) = 2.62[0.03] \text{ ARCH}(12) = 2.15[0.02]$$

where EC is the error correction term, figures in parentheses denote t-statistic values, whereas figures in parentheses denote *p* values. An asterisk indicates significance at 5%. CAP is the dummy variable capturing the restructuring conditions of the CAP policy in May 1992. For the empirical purposes of this section, an 8-lag VAR model was used. The lag selection criteria were based on Akaike's information criterion and Schwarz's information criterion. However, only the variables that turned out to be significant are reported. The dominant features of the estimated model are:

1. Relative food price adjustment to deviations from disequilibria is rather fast, i.e., the estimated speed of adjustment parameter is –0.42.
2. The short-run effect of real public deficit is rather strong. The same also holds for per-capita income, whereas the short-run effect of real money balances as well as of the real exchange rate is relatively low. Fiscal policies seem to exert a more powerful effect on relative food prices than policies based on the monetary spectrum. In other words, fiscal policy seems to feed overall inflation trends in general, and food inflation, which is a major component of overall inflation, in particular. The root cause of such link seems to be the reckless deficit spending that fiscal authorities have resorted to over the period under study.
3. Finally, the CAP restructuring event continues to exert a negative impact on relative food prices even in the short run. These findings imply that the CAP reform led to lower food prices to render them more competitive in the internal and world market.

The estimated model satisfies certain diagnostics such as absence of serial correlation (LM), absence of misspecification (RESET), presence of normality (NO), and absence of heteroscedasticity (HE). However, it suffers from the presence of ARCH effects, even at the first lag. The ARCH tests indicate the presence of volatility clustering in relative food prices. They suggest the use of the GARCH methodology as the appropriate approach to generate both consistent estimates of the mean equation and to describe the evolution of the variance of relative food prices.

GARCH and GARCH-X Estimates

On the basis of parsimony criteria, GARCH models are considered as a special case of an ARMA process (Tsay, 1987). Therefore, through a Box-Jenkins methodological procedure, a GARCH(1,1) model exhibits the best fit. Higher-order GARCH formulations added no significant improvements in goodness of fit. The results yield the following estimates:

$$\begin{aligned}
\Delta pp = {} & \underset{(13.2)^*}{0.391\,\Delta pp(-4)} + \underset{(2.49)^*}{0.058\,\Delta pp(-5)} + \underset{(3.34)^*}{0.108\,\Delta pp(-8)} + \underset{(13.4)^*}{0.415\,\Delta RDEFY(-1)} - \\
& \underset{(3.86)^*}{0.0189\,\Delta RDEFY(-2)} + \underset{(4.28)^*}{0.0161\,\Delta RDEFY(-3)} - \underset{(-5.89)^*}{0.205\Delta rm(-4)} + \underset{(3.27)^*}{0.114\Delta re(-8)} + \underset{(2.85)^*}{0.043\Delta ypop(-3)} \\
& + \underset{(6.28)^*}{0.074\,\Delta ypop(-4)} + \underset{(14.1)^*}{0.239\,\Delta ypop(-6)} + \underset{(11.2)^*}{0.166\,\Delta ypop(-7)} + \underset{(6.34)^*}{0.098\,\Delta ypop(-8)} - \underset{(-7.76)^*}{0.107\,EC(-1)} \\
& \underset{(-6.16)^*}{-0.29\,CAP}
\end{aligned}$$

$$h_t = \underset{(13.4)^*}{0.00357} + \underset{(7.88)^*}{0.506 h_{t-i}} + \underset{(10.6)^*}{0.232\,\xi_{t-i}^{\,2}}$$

Function value (the log likelihood) = 1377.17 with ξ_t being the residuals from the EC model and figures in parentheses denoting t-statistics. Similarly, a GARCH-X(1,1) model is identified, which yields the following estimates:

$$\begin{aligned}
\Delta pp = {} & \underset{(2.56)^*}{0.306\,\Delta pp(-4)} + \underset{(3.32)^*}{0.123\,\Delta pp(-5)} + \underset{(2.18)^*}{0.045\,\Delta pp(-8)} + \underset{(4.96)^*}{0.424\,\Delta RDEFY(-1)} - \underset{(-3.97)^*}{0.0399\,\Delta RDEFY(-2)} \\
& + \underset{(4.59)^*}{0.0298\,\Delta RDEFY(-3)} - \underset{(-5.96)^*}{0.148\,\Delta rm(-4)} + \underset{(5.05)^*}{0.283\Delta re(-8)} + \underset{(3.08)^*}{0.168\Delta ypop(-3)} + \underset{(5.99)^*}{0.056\Delta ypop(-4)} \\
& + \underset{(5.03)^*}{0.279\,\Delta ypop(-6)} + \underset{(3.34)^*}{0.209\,\Delta ypop(-7)} + \underset{(4.89)^*}{0.129\,\Delta ypop(-8)} - \underset{(-5.12)^*}{0.131\,EC(-1)} - \underset{(-3.97)^*}{0.205\,CAP}
\end{aligned}$$

$$h_t = \underset{(11.6)^*}{0.00428} + \underset{(4.51)^*}{0.531 h_{t-i}} + \underset{(7.42)^*}{0.255\,\xi_{t-i}^{\,2}} + \underset{(10.5)^*}{0.0806\,EC_{t-1}^{\,2}}$$

Function value (the log likelihood) = 2459.24 with EC being the deviations (the residuals) from the cointegrating vector. Figures in parentheses denote t-statistics. The coefficient on the EC term is positive and statistically significant, indicating a direct relationship between volatility and short-run deviations. These findings imply that prediction of food prices may become a difficult task as the deviation of that series from macroeconomic factors increases in the short run. In this case, effective policy formulation could be also turn to be a difficult task in the short run. In terms of the Likelihood Function value, the GARCH-X(1,1) model outperforms the standard GARCH(1,1) model, whereas all coefficients in the model satisfy the nonnegativity condition. Finally, another important result is that although the persistence measure remains less than one, it is higher with the inclusion of the EC term, implying that denoting explicitly the macroeconomic shocks leads to higher persistence effects on food price volatility.

Alternatively, we can test the superiority of either model based on nonnested or encompassing tests for volatility models suggested by Chen and Kuan (2002) and Engle and Ng (1993). In particular, we consider an LM test, which is based on a minimal nesting model that considers the $T \times R^2$ form. The idea is to construct a model that encompasses both alternatives; the extended model is an augmented model that has as a special case, either the regular GARCH model or the GARCH-X model. When we test the adequacy of the GARCH model against the adequacy of the GARCH-X model and vice versa, the *p* value in the former test turns out to be 0.02, which rejects the null hypothesis that favors the GARCH model, whereas the same test yields a *p* value of 0.34, which accepts the null hypothesis and supports the GARCH-X model. Thus, the GARCH-X model fits the data substantially better than the regular GARCH model.

A Robustness Test

The goal of this subsection is to quantify food price volatility with a different type of the GARCH model, i.e., with the Exponential GARCH model (EGARCH) or the asymmetric GARCH model, to capture certain characteristics of food prices such as an alternative measure of persistence and asymmetric effects. This model, suggested by Nelson (1991), can successfully model asymmetric impacts of good news and bad news on food price volatility with high levels of accuracy. According to this model, the natural logarithm of the conditional variance is allowed to vary over time as a function of the lagged error terms rather than lagged squared errors. This model is more advantageous than the original GARCH model because, first, it allows for the asymmetry in responsiveness of food price volatility to the sign of shocks to food prices and, second, does not impose the nonnegativity constraints on parameters. Although this model has been extensively used to quantify volatility in various money, financial, and exchange rate variables, to our knowledge, it has not been examined yet to examine food price volatility. The EGARCH(1,1) model can be written as:

$$\ln h_t^2 = \omega + \alpha|\xi_{t-1}/h_{t-1}| + \gamma\xi_{t-1}/h_{t-1} + \beta \ln h_{t-1}^2 \tag{4}$$

The exponential nature of the EGARCH ensures that the conditional variance is always positive even if the parameter values are negative; thus, there is no need for parameter restrictions to impose nonnegativity. The parameter γ captures the asymmetric effect, whereas the parameter β captures persistence. The model is estimated using the robust method of Bollerslev-Wooldridge's quasimaximum likelihood estimator assuming the Gaussian standard normal distribution. The results yield:

$$\begin{aligned}
\Delta pp = {} & \underset{(3.05)^*}{0.287\,\Delta pp(-2)} + \underset{(3.61)^*}{0.146\,\Delta pp(-3)} + \underset{(3.44)^*}{0.102\,\Delta pp(-5)} + \underset{(3.87)^*}{0.406\,\Delta RDEFY(-1)} - \underset{(-4.16)^*}{0.136\,\Delta RDEFY(-1)} \\
& + \underset{(4.16)^*}{0.095\,\Delta RDEFY(-2)} - \underset{(-4.67)^*}{0.155\,\Delta rm(-2)} + \underset{(4.67)^*}{0.261\Delta re(-6)} + \underset{(3.58)^*}{0.179\Delta ypop(-2)} + \underset{(5.47)^*}{0.135\,\Delta ypop(-2)} \\
& + \underset{(5.53)}{0.292\,\Delta ypop(-4)} + \underset{(3.62)^*}{0.227\,\Delta ypop(-3)} + \underset{(4.36)^*}{0.157\,\Delta ypop(-5)} - \underset{(-4.68)^*}{0.157\,EC(-1)} - \underset{(-4.26)^*}{0.245\,CAP} \\
h_t = {} & \underset{(10.4)^*}{0.0258} + \underset{(3.25)^*}{0.248|\xi_{t-1}/h_{t-1}|} + \underset{(5.62)^*}{0.176\,\xi_{t-1}/h_{t-1}} + \underset{(7.36)^*}{0.753\,h_{t-1}^2}
\end{aligned}$$

Function value (the log likelihood) = 2125.73

The results report that the persistence in volatility is higher than the original GARCH model but lower than its counterpart in the GARCH-X model case. Moreover, the asymmetric effect is positive and statistically significant, indicating the presence of an asymmetric impact of good and bad news on food price volatility. In particular, these results suggest that unanticipated increases in food prices lead to food price volatility increases more than in the case of unanticipated decreases in food prices.

Conclusions and Policy Implications

This article investigated the behavior relative food price volatility and the potential effects of short-run deviations between relative food prices and specific macroeconomic factors on food price volatility in Greece. The empirical analysis used the methodology of GARCH and GARCH-X models. Short-run deviations were proxied by the EC term from the cointegration relationship between relative food prices and certain macroeconomic variables such as real money balances, real per-capita income, the real exchange rate, and the real deficit-to-income ratio. The results from a GARCH-X(1,1) model showed that a significant and positive effect is imposed by those deviations on the volatility of relative food prices. Moreover, although persistence remains less than one, the inclusion of macroeconomic shocks gets them closer to permanency. It remains an avenue for future research to examine whether similar results could be confirmed across different country groups.

The implications of these findings are quite critical for the future course of food prices because it relied on the course of the macroeconomic environment, especially now with the current problematic fiscal position of the Greek economy playing the dominant role in the country's macroeconomic environment. An increase in price volatility implies greater uncertainty about future prices, because the range in which prices might lie in the future becomes wider. As a result, producers and consumers can be affected by increased price volatility, because it augments the uncertainty and the risk in the market. Increased price volatility can reduce the accuracy of producers' and consumers' forecasts of future food prices, thereby causing welfare losses to both producers and consumers of food commodities. It is also crucial for policymakers to be aware of the degree of price volatility so as to be able to adopt appropriate hedging strategies. Thus, from a policy standpoint, the results are important, because once the participants receive a signal that the food market is too volatile, it might lead them to call for increased government intervention in the allocation of investment resources and this could not be a welfare improvement factor. However, the global economic crisis may complicate the resolution of many of these macroeconomic factors such as public deficits and public debts. Thus, tighter credit and fiscal positions will make it more difficult to finance the massive investments needed in the agricultural sector to enhance food security, especial for low-income households.

In addition to these recommendations, additional certain actions must be also implemented as to alleviate the impact of the macroeconomic environment on food price uncertainty such as encouragement of financial institutions to expand operations rapidly to improve access of farmers to credit and the encouragement by the Greek authorities of significant increases in investments in adaptive research and technology dissemination (though the current situation in the Greek economy makes the feasibility of such efforts extremely doubtful).

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