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Has world cereal market instability increased?

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Abstract

This paper argues that world cereal prices are good indicators of the state of world cereal markets. It tests whether real prices for wheat, rice and maize exhibit deterministic or stochastic trends, concluding that the underlying trends are most likely deterministic. After appropriate deterministic detrending, the paper tests for increased variance of the residuals of real cereal prices, finding that there does not appear to be evidence for increased year-to-year price instability. Tests for increased intra-year price variability also reject the hypothesis of increased variability. Discussion of recent changes in the pattern of world cereal production and trade suggests that these conclusions are not contrary to observed trends.

Keywords: Cereals; Instability; Prices

Introduction

Price and more generally market instability in cereal markets have been the concern of private producers and consumers as well as governments for a long time. The reason is that market volatility, especially unforeseen price variations in response to exogenous or endogenous shocks, can lead to sudden and large income transfers among various market participants. A relatively new focus of the more recent concerns is whether the nature of world market instability in cereals has changed over the last decade. There are several reasons which could suggest that possible changes are occurring in market instability. For instance, world trade has become more liberalized, and the role of governments in cereal market interventions has lessened. Several major producing countries (China, the Former Soviet Union and Eastern Europe) are undergoing significant economic restructuring. Technological changes

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in production might have induced more unstable cereal yields. The communications revolution has led to increased integration of regional markets. The purpose of this paper is to explore the changing nature of world cereal market instability.

It is useful before embarking on empirical analysis to settle some conceptual issues concerning instability. The first relates to the appropriate index with which to measure instability in a commodity and particularly a cereal market. It is well recognized that cereal markets are well developed, and integrated, in the sense that there are "focal" markets that provide lead signals for the world-wide state of the market. For instance, for wheat the Chicago market is probably the most important one in the world. This does not mean that there are no other appropriate indicator prices. In the case of wheat, for instance, the US Gulf export price is a lead price for the world wheat export market. Similarly in every country there are indicator prices for most commodities. These prices are in general related to the focal prices in the sense that they tend to have similar movements in response to important events (Mundlak and Larson, 1992). Of course, localized events can lead to temporary deviations of prices from the indicator prices. The above discussion brings into focus the fact that generally it is the price (spot and/or forward) that is regarded as the key signal for the state of a commodity market.

If prices are accepted as the summary measures of the supply/demand situation of a commodity market, then the question arises concerning the appropriate description of market instability. An easily computed measure is the period-to-period (daily, monthly or yearly) variation in price. This might be appropriate if all the change in price from period to period is unanticipated and hence unknown. This is not, however, true, as for several periods price changes are expected in response to known market outcomes. For instance in a closed cereal market, it is expected that the price in the early harvest period will be lower than the price in the late part of the crop year. This type of seasonal price variation in the course of the year cannot be regarded as part of price instability. One might attempt to account for the known influences on periodic price changes, in order to isolate the truly random unforeseen events, but this is not always easy, as these factors can change over time.

A related way to measure price instability is to try to construct a model of the underlying "trend" price. Instability can then be defined as the deviation of the observed price from the trend. The problem is, of course, that it is not easy to construct the trend. A trend for a period t should be defined as the market expectation of the price in period t, based on information up to some previous period (say s periods before). Since information at time $t-s$ will be a function of t and s, namely will be changing from period to period, the trend price itself when defined in this fashion will be an unstable variable, and in addition will not be easily measurable.

The above point is subtle and important. Consider, for instance, the problem of defining a trend line for period t, given information up to period $t-1$. This trend line, if it is to summarize all information up to time $t-1$, should include the realized values of market-related variables up to time $t-1$, such as the price at time $t-1$. On the other hand, if the trend is to include only "long-run" and slowly changing events, such as technological changes, etc., then it should not include the realized values of market variables at time $t-1$ and earlier, but rather values of other related variables that are not necessarily changing from period to period. Such a variable could, for instance, be a simple time trend. It is clear that the magnitude of instability will be assessed in a different fashion under the alternative definitions, and hence one should be careful to state clearly the assumptions involved in defining the “trend”, or underlying “expectation”, departures from which constitute instability. The modern theory of cointegration has made several efforts at defining and estimating “stochastic trends” but it appears that there is no consensus on the best practice (see the recent debate on this in the *Economic Journal*, articles by Granger, 1997; Pesaran, 1997; Harvey, 1997).

In Trends in yearly cereal prices tests are conducted to understand the nature of the underlying trends in cereal prices. Then, in Instability in annual cereal prices, tests for changes in inter-year price instability in cereal markets are elaborated. Intra-year price variability discusses the issue of intra-year cereal price instability. Next, The changing world pattern of cereal production and trade is discussed in an effort to relate it to the empirical instability findings. The results are summarized in the Conclusions.

Trends in yearly cereal prices

The issue in this section is to explore the types of trend exhibited by the yearly prices in world cereal markets. The data are the following. For wheat, the price for US No. 2 Hard Winter Ordinary fob Gulf is used. For coarse grains, the price for maize US No. 2 Yellow fob Gulf, and for rice the price for white Thai 5% broken fob Bangkok were utilized. All prices are in US$ per metric ton (mt). The US monthly consumer price deflator (which averages 100 for year 1983) was used to deflate the monthly data.

From the deflated monthly data two types of yearly simple average were computed. One was for calendar years, and the second was for July–June crop years. A priori, one would want to work with crop-year prices, as these seem to be more relevant for the bulk of world production and marketing of the cereals wheat and maize. The calendar-year prices are highly correlated with the crop-year prices. The various initial tests were conducted on both crop-year as well as calendar-year prices, and on both nominal and real prices, albeit the real ones are those of main interest. The subsequent analysis focuses on deflated crop-year data.

As indicated earlier, it is deviations from appropriately defined trends that should be analyzed to examine issues of instability. It is important as a first step to investigate whether the series are trend stationary (TS) or difference stationary (DS). Trend stationary series are those that can be described as the sum of a deterministic time trend and a stationary process (the latter being a process whose mean and variance do not vary with time). Difference stationary, or unit root processes, are those whose first difference is a stationary process. The major difference between the two, as has been outlined in several recent econometric textbooks, is that in TS processes disturbances tend to have temporary effects, namely do not lead to permanent shifts of the process, while in DS processes any shocks tend to leave permanent effects on the process (see, e.g., Hamilton, 1994, Ch. 15; Enders, 1995, Ch. 8). This is, of course, quite important to know for the world prices of cereals since, if they can be characterized by DS processes, any temporary shocks to prices would have permanent effects.

To test whether the annual world prices of cereals are characterized by TS or DS processes, a procedure outlined in Enders (1995, pp. 256–258) is utilized. Denote an arbitrary time series by x_t, and by Δx_t the first difference of the series ($\Delta x_t = x_t - x_{t-1}$). The procedure consists of the following steps.

First, estimate by least squares (LS) an equation of the form:

$$\Delta x_t = \alpha_0 + \alpha_2 t + \gamma x_{t-1} + \sum_{i}^{n} \beta_i \Delta x_{t-i} + \varepsilon_t, \tag{1}$$

where t denotes a linear time trend, ε_t denotes the error term, and the Greek letters denote coefficients. The size of the maximum lag n is determined by simple t-tests on the coefficients β_i.

Second, perform an augmented Dickey–Fuller test (ADF) on whether the parameter γ that multiplies the undifferenced term is zero. Such tests have become common in recent years in the context of the so-called unit root revolution in econometrics (for detailed descriptions of the methods, see Hamilton, 1994; Enders, 1995). If the test shows that the parameter is non-zero, then stop and conclude that the series does not contain a unit root, and hence is TS. If the test shows that γ is zero, then proceed sequentially to first test whether the coefficient of the trend term is zero, based on some specialized tests, and if it is, re-estimate the equation without the trend term, redoing the ADF test for testing the zero value of γ. Then test whether the constant term is zero and, if it is, redo the ADF test in a re-estimated equation. At each step if it is found that γ is not zero, then one stops and concludes that the series does not contain a unit root. Otherwise one accepts the hypothesis of a unit root.

Table 1 exhibits the results of this type of test on all the initially estimated world price series for cereals, namely all 12 series. For all series, the maximum value of n in Eq. (1) was found to be equal to 1, and for most series the exclusion of the trend was not necessary. For almost all series, nominal or real, calendar- or crop-year-based, the hypothesis of a unit root seems to be strongly rejected. From all the 12 series tested the only one for which the hypothesis of a unit root could not be rejected is the series for the real price of maize computed in a calendar manner. The results of similar tests where the corresponding series are the natural logarithms of the relevant prices are almost identical to those of Table 1.

As considerable cereal market instability was exhibited in the period 1973–1975, it might be hypothesized that the trend behavior of prices is different before and after this period. If years prior to 1976 are omitted from the series and the unit root tests are redone, as outlined above, then the unit root hypothesis is rejected again for the real crop-year annual prices of maize and rice, but is not rejected for wheat. If the tests on the logarithms of the real crop-year prices are redone, then the unit root hypothesis cannot be rejected for any of the series.

Table 1
Results of testing for unit roots in annual world cereal price series (the data are for the period 1970–1996. The entries in the table indicate whether the null hypothesis of a unit root is accepted and the degree of confidence[a,b])

Commodity	Type of annual price utilized			
	Calendar year		Crop year	
	Nominal	Real	Nominal	Real
Wheat	No** (1) (tr)	No*** (1) (tr)	No** (1) (tr)	No*** (1) (tr)
Maize	No* (1) (c)	Yes (1) (tr)	No** (1) (c)	No*** (1) (tr)
Rice	No** (1) (tr)	No*** (1) (tr)	No** (1) (tr)	No*** (1) (tr)

[a] * indicates rejection of the unit root hypothesis at 10% significance level; ** indicates rejection of the unit root hypothesis at 5% significance level; *** indicates rejection of the unit root hypothesis at 1% significance level.

[b] The number in the first set of parentheses after each entry denotes the number of significant lags incorporated in the test regression. The second set of parentheses denotes whether a constant drift plus trend (tr) was included in the test regression, whether simply a constant drift (c) was included, or whether no drift and no trend (n) was included.

Of course, the data in the series are not long enough to be able to discern statistical stationary patterns that should normally be identified from long-time series, and this makes the whole analysis of unit roots somewhat suspect. In fact, one of the major criticisms of unit root tests is that they are have low power against stationary models with roots close to unity (Rudebusch, 1993), and are not robust against alternative specifications that include trends with breaks (Hendry and Neale, 1991). Leon and Soto (1995) found in their examination of long annual real commodity price series (1900–1992) that while standard unit root tests could not reject the hypothesis of unit roots for almost all series, once the possibilities of breaks was admitted, the unit root hypothesis was rejected for most series. In their analysis, using a test that is robust to structural breaks, they found that the long-run behavior of maize, wheat and rice price series did not contain a unit root, and they could best be described by TS processes. This is similar to the conclusion reached above, with admittedly shorter series. Given, however, the coincidence of the results on both short- as well as long-time series, it will be assumed in the sequel that the cereal price series that are analyzed (using crop-year deflated data) are characterized by TS processes.

The admission of TS processes for the price series does not imply the lack of structural breaks. An attempt was made to test the trends for structural breaks. This is important, because one might mistake a structural break in a trend for an increase in the variance of the series. The procedure utilized was the following. First a linear trend was fitted to the data (in absolute or logarithmic form). The coefficients of this regression were tested for stability using a variety of tests. The tests included the plot of recursive residuals, the CUSUM test, the CUSUM of squares test and the plot of recursive coefficients.[1] If these tests suggest that there is a break in the coefficients of the trend in some year, then a Chow breakpoint test was performed for that year. This test involved dividing the period into two subperiods indicated by the breakpoint tests, estimating separate trends for each of the subperiods, and testing whether the coefficients from each of the regressions are the same.

The results of the tests were inconclusive, basically because of lack of sufficient degrees of freedom, namely number of years of observation. This was pointed out also by Leon and Soto (1995), who used much longer annual series and even then had difficulty finding an appropriate test, the standard tests utilized here being very weak. In general, it appears that there is some type of structural break after 1977, and the possibility of another after 1983, but again the degrees of freedom are too few for any conclusive results.

Instability in annual cereal prices

Given the above results, it was decided to detrend the (real) price series by simple linear trends in the levels. Table 2 indicates the regression output from the simple linear time trends and the value of the Breusch–Godfrey serial correlation specification test. It is quite obvious that all trends are negative and significant. It is also quite evident that there exists serial correlation in the residuals of the trend regressions. Given this observation, standard Box–Jenkins identification techniques

Table 2
Results of linear trend regressions[a] on the annual real crop-year prices of cereals (period of estimation is 1971–96)

	Wheat	Maize	Rice
Constant	262.240***	211.150***	594.851***
Linear time trend	−6.519***	−5.642***	−17.326***
Correlation R^2	0.467	0.634	0.382
Durbin Watson statistic	0.824	0.794	1.023
Breusch–Godfrey serial Correlation LM test[b]	8.595***	11.471***	5.593**

[a] * denotes significance at the 10% level; ** denotes significance at the 5% level; *** denotes significance at the 1% level.

[b] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of no serial correlation of residuals is rejected.

[1] In recursive least squares the equation is estimated repeatedly, using increasingly larger subsamples of the data, starting with the minimum possible number of observations. From each regression, the coefficient estimates are used to produce a one-period-ahead forecast. The difference between the actual value of the series and this one-period forecast is the recursive residual. The CUSUM test involves plotting the cumulative sum of recursive residuals. The CUSUM of squares involves the plot of the cumulative sum of squared residuals. Parameter instability is indicated when the cumulative sums go outside a plotted significance area.

were, therefore, used to specify appropriate autoregressive moving average (ARMA) models for the residuals of the trend regressions.

Table 3 indicates the results of linear trend regressions of world crop prices allowing for ARMA specification of the residuals. The models are of the following general type:

$$y_t = \alpha + \gamma t + u_t, \tag{2}$$

where y is the price in year t, u is the error in year t that follows an ARMA(p, q) process of the type

$$u_t + \sum_i^p a_i u_{t-i} = a_0 + \sum_i^q b_i \varepsilon_{t-i}. \tag{3}$$

The basic criterion for specifying the models is that they are as low order as possible, and that the estimated models exhibit no more serial correlation. Furthermore, tests of normality indicate for all models that the residuals are normal. The table indicates the best models fitted, the values of the fitted parameters a_i and b_i with their levels of significance, and three residual tests. The first is the Breusch–Godfrey test for serial correlation, and the other two are tests for heteroscedasticity of residuals. The first of these tests for low-order autoregressive conditional heteroscedasticity (ARCH), while the second tests for general type of heteroscedasticity. The ARCH

Table 3
Results of linear trend regressions[a] on annual real crop-year prices of cereals with ARMA errors (period of estimation is 1971–96)

	Wheat	Maize	Rice
Constant	281.571***	232.107***	691.268***
Linear time trend	−7.353***	−6.707***	−22.479***
AR(1)	0.705***	0.768***	0.511**
AR(2)	−0.578***	−0.508***	−0.402**
MA(1)	−0.399***		
MA(2)	0.984***		
Breusch–Godfrey serial Correlation LM test[b]	0.666	0.757	1.685
ARCH LM test statistic[c]	0.992 (1)	1.959 (1)	0.425 (1)
White's heteroscedasticity test[d]	0.466	0.834	3.762**

[a] * denotes significance at the 10% level; ** denotes significance at the 5% level; *** denotes significance at the 1% level.

[b] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of no serial correlation of residuals is rejected.

[c] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of ARCH type of heteroscedasticity of residuals is rejected. The lag included in the ARCH regression is indicated next to the value of the test.

[d] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of no heteroscedasticity of residuals is rejected.

models originally proposed by Engle (1982) assume that the conditional variance of a variable, namely the variance of a one-period-ahead forecast, given values of the variable in previous periods, is not constant, and instead depends on recent shocks. This type of model appears relevant for the study of world price instability as it assumes that recent events lead to some type of temporary instability that eventually dies out.

The results indicate that, except for rice, the fitted ARMA models do not exhibit heteroscedasticity, meaning that the conditional variance of the series, namely the variance of a year's price, given information of previous years, does not appear to vary. This is also the case for rice when the ARCH test is done, but it appears that with rice there is some other type of heteroscedasticity in the data. This is indicated by the significant value of the White *F*-test.

The tests were redone using the logarithms of the annual prices. The conclusions were the same, and in fact even stronger in the sense that even for rice the White test for heteroscedasticity did not reject the hypothesis of no heteroscedasticity. Therefore, the conclusion from this empirical examination is that there does not appear to be an increasing degree of inter-year variability in world cereal markets. Recent events do not appear to manifest anything considerably unusual, or much outside the range of normal annual variations.

Intra-year price variability

The next issue that is investigated relates to the degree of intra-year price variability. To this end the following manipulations to the data were done. First, for each real commodity price, and each crop year, the variance of the 12 monthly prices included in the July–June crop year was calculated. This variance was divided by the average crop-year price. The resulting numbers are the coefficients of variation of intra-year prices, and are unitless. These numbers are reasonable measures of the intra-year price variability of a commodity. The subsequent analysis has as objective to investigate whether there are any trends in these coefficients of variation.

Table 4 indicates the results of linear trend regressions in these coefficients of variation. Apart from the trend results, the table includes test statistics for serial correlation, ARCH and heteroscedasticity. The first observation is that for all three commodities the coefficients of the trend regressions are insignificant. Hence, there does not appear to be any tendency for the intra-year price variability to change. In fact, the coefficients of the trend regressions are all negative, albeit non-significant, implying that the tendency, if any, is for a reduction in the intra-year cereal price variability.

The specification tests indicate that there is no serial correlation among intra-year price variabilities. The value of the ARCH and White tests, however, show that there seems to be some heteroscedasticity in the intra-year price variability of wheat, but not for the other two commodities. This means that while there does not appear to be any trend in the intra-year wheat variability, there might be some time variation of the magnitude of the instability, albeit not of the trend type.

Table 4
Results of trend regressions[a] of the coefficients of variation of intra crop-year prices (period of estimation is 1971–96)

	Wheat	Maize	Rice
Constant	0.0938***	0.0849***	0.1155***
Linear time trend	−0.0014	−0.0007	−0.0011
Breusch–Godfrey serial Correlation LM test[b]	0.354	0.469	1.156
ARCH LM test statistic[c]	10.977*** (1)	0.983 (1)	0.007 (1)
White's heteroscedasticity test[d]	10.476***	2.148	0.082

[a] * denotes significance at the 10% level; ** denotes significance at the 5% level; *** denotes significance at the 1% level.

[b] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of no serial correlation of residuals is rejected.

[c] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of ARCH type of heteroscedasticity of residuals is rejected. The lag included in the ARCH regression is indicated next to the value of the test.

[d] This is the value of the *F*-test. One, two or three stars (denoting significance at 10%, 5% or 1%, respectively) indicate that the null hypothesis of no heteroscedasticity of residuals is rejected.

The results, therefore, seem to support the conclusion that there is no increasing trend in the intra-year variability of world cereal prices. This, of course, does not mean that the absolute values of intra-year variations do not change. In fact, in a year of high average prices the monthly variations are expected to also be high, while the opposite is expected in a year of low average prices. This could be explained partly by the fact that years of high prices are normally associated with low volumes of stocks, and hence any news concerning the market developments tend to lead to larger reactions by market participants. The opposite is normally the case in years of low prices. The proper way to analyze volatility, however, is in relative terms, as was done here, by using the coefficient of variation, and it is on the basis of analyzing such coefficients that the conclusion is that there has not been any trend in intra-year price volatility.

The observation, nevertheless, that in periods of high prices even the coefficient of variation seems to increase in some commodities (notably wheat) might suggest that in such periods the wheat demand might become very inelastic, thus leading to this excess seasonal variation. This is a topic that might merit further research.

The changing world pattern of cereal production and trade and consequences for instability

Over the past 30 years in the world wheat market, while total sown area in the world has not changed, the Former Soviet Union (FSU) has diminished its share by more than 10 percentage points, while Asia — mainly China and India — have gained world wheat area shares. In terms of production this pattern is even more pronounced, with China and India doubling their world shares of world wheat production, while the share of the FSU has halved. Over the past three decades world wheat yields have almost doubled. Currently Canada and the US account for 16% of world wheat production, the EU15 for another 16%, the FSU for 13%, China for 18.4%, and India for 10.7%. The major other exporters, Australia and Argentina, account for less than 5% of world wheat production.

World rice area has expanded by 15% in the last three decades, while production has doubled. Area and production shares have not changed by much, with Asian countries accounting for 91% of world production. China alone accounts for 35% of world production, India for 22%, Indonesia for 9%, and Thailand for 4%. Among members of the Organization for Economic Co-operation and Development (OECD), only Japan is a major producer accounting for 2.5% of world production.

The world coarse grain market has not exhibited any major changes in the allocation of production. Total sown area has stayed constant, while production has increased by 60%. Among major producers, currently the US accounts for 28% of world coarse grain production, the EU15 for 11%, FSU for 9.5%, Brazil for 3.7%, China for 14% and India for 3.6%.

World cereal yields have increased steadily over time, but do not seem to exhibit any pattern of increased variability around the trends. This seems to counter those who have suggested that the introduction of some types of modern yield-enhancing technology (such as for instance hybrid seeds) has increased yield variability. Of course, testing the aggregate world cereal yields might hide significant differences among yields by region or country.

As the world cereal markets are not characterized by free trade, the important thing for world instability is the degree to which various countries allow their domestic production disturbances to be transmitted to the international markets. An appropriate measure of this is the transmission coefficient. Transmission coefficients (for a definition and first results, see Blandford, 1983) measure the proportion of domestic production disturbances that are transmitted to world markets, by variations in net import volumes. These coefficients are influenced by domestic short-run demand and supply price elasticities, by the behavior of domestic stockholders, and most importantly by government policies regarding domestic price stability and stockholding. They are expected to be negative and smaller than one in absolute value.

Such coefficients were estimated for cereals by Sarris (1998) for all the major countries and country aggregates that constitute the world cereals markets. The results showed that for most countries and regions the transmission coefficients were in the expected range and significant. On the basis of the estimations, transmission coefficients were assigned to all countries and regions producing cereals. Subsequently the magnitudes of the production variabilities transmitted to the world by each country and region and at 5 year intervals were estimated as follows. The estimated coefficients of variation of production for each country were multiplied by the average production of cereals for each 5 year period, and finally multiplied by the estimated transmission coefficients. The sum of these numbers divided by the average volume of world imports can be taken as a summary measure of world production variability transmitted to the world, and thus contributing to instability. Table 5 presents these summary results.

It can be seen from Table 5 that for wheat, after a period of high instability in the pre-1980 period, the decade of 1980–90 was characterized by a lower overall degree of instability. This seems to have been reversed again in the 1990s, with the index of instability rising again to levels similar to those of the turbulent 1970s. For rice the picture seems to be quite different, with the degree of instability in the 1990s being at the lowest level compared with all the previous periods. For coarse grains, the instability situation in the 1990s appears to be even worse than in the decade of the 1970s.

In interpreting the above results it must be kept in mind that the implicit assumption made in deriving the summary indicators is that the production variations are uncorrelated between countries or regions. This is not correct, as weather patterns tend to have cross-country and cross-regional influences. Nevertheless, it can be taken as a first step in the analysis of production variability transmitted to the world.

If production variability, as exhibited in transmitted variations, was the only factor

Table 5
Cereal production variations transmitted to world markets (for production variations and imports, the units are thousands of metric tons)

	1965–70	1971–75	1976–80	1981–85	1986–90	1991–96
Wheat						
Total transmitted variation	13,554	15,485	19,001	21,458	23,285	23,232
Average world imports	54,541	67,493	81,594	107,900	109,613	100,814
Ratio transmissions to world imports (%)	24.85	22.94	23.29	19.89	21.24	23.04
Rice						
Total transmitted variation	1200	1415	1644	1817	2021	2384
Average world imports	7535	7839	10,092	11,386	11,853	17,098
Ratio transmissions to world imports (%)	15.93	18.05	16.29	15.96	17.05	13.95
Coarse grains						
Total transmitted variation	10,075	12,509	13,979	19,868	19,899	20,964
Average world imports	45,438	70,032	10,0404	111,353	107,456	91,120
Ratio transmissions to world imports (%)	22.17	17.86	13.92	17.84	18.52	23.01

Source: Sarris (1998).

affecting world cereal markets, then one would expect, on the basis of the above results, that world prices should be more unstable in the 1990s in the wheat and coarse grain markets, compared with the earlier parts of the decade, while for rice the picture would be one of lower degree of instability. The price analysis, however, did not support such a conclusion, although there were not enough degrees of freedom for a statistically significant statement.

Production variations, however, are only part of the factors affecting world cereal markets. In fact, in his simulations of world cereal markets Mitchell (1987) found that about 50% of the world price variations were due to macroeconomic and other exogenous shocks (mainly the oil price shock), another 25% was due to agricultural policy shocks, and only 20–30% of the variation was due to unexplained factors, presumably weather and model mis-specifications. Of course, his analysis was conditioned to a large extent by the two oil shocks and the macro events that accompanied them. A more recent analysis by Sarris (1990), despite its aggregated nature, also found that a major part of the world cereal price instability was due to non-production factors. Nevertheless, the importance of production and agricultural policy variability, even for the tumultuous period that Mitchell examined, accounted for 40–50% of the world price variations.

A caveat of the analysis and conclusions is that it was assumed that the transmission behavior of countries is the same over the whole period of the last three decades. This is not necessarily a good assumption, as over this period there have been significant policy changes that could have affected behavior. It was not possible to test for changing behavior, however, with the limited time-series data available.

Nevertheless, the above caveat does not invalidate the general conclusions reached concerning the transmission of production fluctuations. If transmission has increased in the 1980s, as one would expect given the general tendency for market opening of a number of economies, this would imply larger proportions of domestic variations that are transmitted internationally. This would increase the measures of instability for all commodities that were exhibited in Table 5. With the exception of rice, this would not affect the overall conclusion that instability appears to have increased in the 1990s in the wheat and coarse grain markets, and in fact would make the conclusion stronger. For rice it could reverse the conclusion of decreasing instability.

Analysis of cereal stocks (reported in Sarris, 1998) suggests that the recent period has seen a decline in the geographical concentration of world cereal stocks. In other words, more countries and regions now hold large end-of-season stocks, compared with the early 1980s. This is a trend that would tend to mitigate world price instability. On the other hand, there has been a declining trend in the ratios of stocks to apparent consumption, and a tendency for a larger share of stocks to be held by the private sector, developments that should tend to destabilize cereal markets.

The overall conclusion of the analysis is that while there are factors that would tend to increase the world instability of cereal markets, there are other counteracting factors that would tend to diminish it. It seems, tentatively, that there does not seem to be a general trend toward increasing world cereal market instability.

Conclusions

Analysis of the changing pattern of world cereal instability has led to several conclusions. First, it was demonstrated that annual cereal prices, whether on a calendar or crop-year basis, nominal or real, seem to be described best by trend stationary (TS) time-series processes. This implies that any temporary shocks to the world cereal markets do not leave permanent effects on prices. This is an important conclusion, and one that merits further investigation with longer time series. This conclusion does not negate the possibility of structural breaks in the cereal markets, but it must be realized that structural breaks are once-and-for-all events that have permanence. Subsequent random shocks are not expected to lead to any permanent changes. The analysis of Leon and Soto (1995) is a good methodological step in the right direction. A topic for further research is a longer-term analysis of world cereal prices, with the purpose to identify structural breaks in the series, and especially after 1973–75. That period has been considered by many as important for changing the world cereal market scene, but no one has analyzed using modern time series tools the type of structural break that occurred then.

The analysis of inter-year price variability of cereals concluded that there does not appear to be an increasing degree of inter-year variability in world cereal markets. Recent events do not appear to manifest anything considerably unusual, or much outside the range of normal annual variations. It was observed that there appears to be some tendency toward increased volatility in the most recent period (1995–96) but it is difficult on the basis of very few observations to be definitive.

Finally, the analysis of the intra-year price variability concluded that there does not seem to be any tendency for the coefficients of variation of monthly seasonal prices to increase over time, and, if any, the tendency is towards a decline.

The overall answer then to the question of whether recent evidence suggests that the world cereal markets have become more unstable recently is “No”. This, of course, does not answer the next logical question, which is whether these markets have become more stable. Trade liberalization and the opening of several hitherto closed or state-controlled markets would suggest that this should be the case. The econometric tests performed are too weak for a conclusive test of this hypothesis, just as they are weak for the test of the increased instability hypothesis, and clearly more data are needed for a better analysis. However, until the time when such data become available, it is probably reasonable to accept that the structure of world cereal price behavior does not seem to have changed much in the last two decades.

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