PN-AAY-459
ISN 52677

A Variance Components Approach to Food Grain Market Integration in Northern Nigeria

Christopher L. Delgado

Reprinted from the
American Journal of Agricultural Economics
Vol. 68, No. 4, November 1986

INTERNATIONAL
FOOD
POLICY
RESEARCH
INSTITUTE

1776 Massachusetts Avenue, N.W.
Washington, D.C. 20036 U.S.A.

Reprint No. 101

Reprinted from
AMERICAN JOURNAL OF AGRICULTURAL ECONOMICS
Vol. 68, No. 4, November 1986

A Variance Components Approach to Food Grain Market Integration in Northern Nigeria

Christopher L. Delgado

A variance components methodology is developed for joint tests on a sample of time series of prices for seasonal differences in the price integration of markets. The approach requires a statistically adequate number of observations for each market within seasons characterized by constancy of transactions costs among markets. The model is applied to eighteen months of weekly grain prices for twenty-two villages in northern Nigeria. Results suggest that markets are not well integrated in the six months covering the harvest period. Implications are drawn for research on market performance in the region.

Key words: grain prices, market integration, Nigeria, seasons, variance components.

The integration of food grain markets is central to the direction and success of agricultural policies in the West African Semi-Arid Tropics (WASAT). A well-integrated market system is essential to household food security in both food deficit rural areas and those witnessing a rise in the relative importance of nonfood cash cropping. It is also key to the sustained success of extension of new food production technologies, which might otherwise flood stagnant local markets with unsalable surpluses. It can also be shown that the degree of food grain market integration determines whether equity-oriented production policies in less favored food-producing areas should be oriented toward food or some alternative activity (Delgado).

Despite the importance of the subject to current policy debates, there has been remarkably little progress in the WASAT in generating the data and models necessary to explore market integration issues further. In a comprehensive survey of the literature, Harriss shows that most work in the region has relied upon computing bivariate correlation coefficients of market price time series, frequently on undetrended data. With few exceptions, analysts in the region must make do with time series for prices of major grains in different locations. Complementary data on transport and storage costs, their variation across locations and time periods, and aggregate flow through the hierarchy of markets at any given time period are very difficult and expensive to obtain.

This article starts from the premise that in most cases analysts will have to continue to rely upon the examination of the covariance of time series on food prices to assess the existence of market integration. This raises two interrelated problems. First, previous analyses of market integration in the region based on bivariate correlation coefficients have tended to find numerically low price correlations despite other evidence that suggests competitive and rational behavior by a large number of market actors. Second, the statistical interpretation of bivariate correlation coefficients in this context is problematic. There is no obvious way to compare and synthesize results once more than two markets

Christopher L. Delgado is co-coordinator for African Research at the International Food Policy Research Institute, Washington, D.C.

Survey data were supplied by the Agricultural Projects Monitoring Evaluation and Planning Unit, Kaduna, Nigeria. Financial support was received from the International Development Research Center, Canada.

An earlier draft, stressing implications for regional equity concerns, was presented at the International Crops Research Institute for the Semi-Arid Tropics, Hyderabad, India, in October 1983.

The comments of Marc Nerlove, John Strauss, and an anonymous reviewer are gratefully acknowledged, as is the research assistance of Hung Nguyen. The author accepts sole responsibility for the views and remaining errors herein.

Review was coordinated by Hans P. Binswanger, associate editor.

Copyright 1986 American Agricultural Economics Association

are involved, since each pair is considered in isolation. Recent methodological work on market integration is still based on pair-wise comparison of price series, albeit with greater attention to detrending and weighting of observations (Granger, Haugh, Gupta and Mueller). Yet, what matters in the present instance is that the system of markets be integrated and not that a particular pair be integrated at any one time.

This article will examine these problems in a case study of northern Nigeria, the zone of the WASAT where food grain market integration has been most intensively studied. Jones' seminal work on seven crops countrywide using thirteen years of undetrended monthly data found that "of 4,836 [bivariate correlation] coefficients only 19 are as great as .90, and 424 are zero or negative" (p. 110). Hays found weak correlations for eleven years of undetrended monthly wholesale price data for fifteen cities in the northern part of the country. Harriss sees these "weak" correlations as useful evidence of the absence of integration, even though the reverse argument would presumably not be accepted for "high" correlations.

Yet the very few studies that have included non-price variables in the quantitative analyses of market integration in the region have generally suspported the conventional wisdom that grain markets function smoothly and in a competitive fashion, given the constraints imposed on individual agents. In particular, Hays and McCoy found that:

> Price differentials exceeding positive price spreads among markets did not result from planned manipulation under monopolistic or monopsonistic conditions but rather from imperfections inherent in the system, which made effective arbitrage in response to spatial price differentials difficult. Intermarket price spreads (both greater than, and less than, transfer costs) exist because erratic supply increases risks in intermarket arbitrage, dissemination of information on prices and supply in the various markets is inadequate, and traders taking part in arbitrage within these markets are non-specialized. (page 191)

Despite the difficult conditions facing traders, the consensus is that grain markets generally operate efficiently in the region (Hays, Hays and McCoy, Jones, Ejiga).

The apparent contradiction between informed belief and quantitative results of correlation analysis will be investigated using a new primary data set of village level weekly grain prices. The model proposed evaluates jointly the information from a series of markets, to assess the integration of the system. Some surprising conclusions for both methodology and for market integration in northern Nigeria emerge.

The approach is to decompose the variance of food grain prices into components, following the classic work by Balestra and Nerlove on the demand for natural gas. The model is designed to handle weekly food grain price data from a sample of village markets in a relatively small region subject to well-known seasonal influences.

The model of price for a specific crop can be written,

$$(1) \qquad p_{it}^{(s)} = m^{(s)} + v_i^{(s)} + u_t^{(s)} + z_{it}^{(s)}.$$

Here the superscripts and subscripts are as follows: $s = 1, \ldots,$ number of seasons, where a season is defined as a continuous number of weeks within the year for which it is reasonable to make two important assumptions. First, the variance of prices for a given crop in a given village is constant over the season. Second, transport and transactions costs for marketing a given crop between any two markets within the system are constant for the two markets concerned, subject to a random disturbance, over the season. As a practical matter, seasons need to contain a sufficient number of weeks to permit statistical estimation of variances. Also, $i = 1, \ldots,$ number of food grain markets (or villages); $t = 1, \ldots,$ number of weeks in season; p_{it} is the price of the grain concerned in market i in week t; m, the mean price of food each season in the sample; v_i, a constant location (village) effect; u_t, a constant weekly time effect; and z_{it}, a stochastic interaction term. For convenience, the superscripts will be dropped, but it should be borne in mind for all that follows that the analysis is done separately for each crop and season.

Thus, the variation across villages and weeks around the sample-wide mean price of a given grain in a given season is divided into two constant deviations and a stochastic term. The deviations from the mean are either constant for a given village (v_i) or for a given week (u_t), but not across both weeks and villages.

The proposed yardstick of integration of a pair of village food grain markets within a given season is that the price spread between them stays approximately constant, subject to random variation either way, over the weeks in the season in question. Put another way, if

the mean price spread between any pair of villages can be reasonably judged random, then it follows that intermarket price differentials are equal over time to a constant, attributable to transactions and transport costs, with deviations from the constant being random noise. Because we assume that transactions and transport costs are in fact constant between any pair of markets within a relatively small region and within a well-defined marketing season, statistical confirmation of the absence of nonrandom divergent price trends over the season between a pair of villages supports the hypothesis that there is a constant relationship between the series. This is, in fact, akin to a bivariate correlation coefficient calculated for the two price series, which measures the degree of linear dependence between them.

Unlike the bivariate correlation coefficient, however, the formulation in equation (1) permits a joint test of the absence of nonintegration between any two pairs of village markets in the sample. This involves showing that after removing a common seasonal trend (u_t) and seasonal mean prices for each village ($m + v_i$), interactions between the residual price terms (z_{it}) across villages are independent; more formally, that

$$E(z_{it}z_{lt}) = 0 \text{ where } i \neq l. \tag{2}$$

If (2) can be shown to hold jointly with a reasonable degree of statistical confidence for all pairs of markets in the sample, then by the same criterion, the system of markets is judged integrated.

A test of (2) above requires a means of estimating the z elements from the price data. This first requires removing long-term trends and constant effects peculiar to a particular location by subtracting mean seasonal prices for each village and crop from the weekly price data. Thus,

$$p^*_{it} = p_{it} - \bar{p}_i, \tag{3}$$

where the notation is consistent with (1) above and the mean price for village i is calculated separately for each crop and season. Therefore,

$$p^*_{it} = u_t + z_{it}. \tag{4}$$

The next step in retrieving the stochastic z_{it} is to eliminate the weekly time effect, u_t, which is conceptually constant across villages but different for each week. Netting this effect out removes the common "seasonal" trend composed of a constant weekly effect common to all villages over the year (but not across years) and thus eliminates spurious correlation of price movements across villages arising from seasonal influences common to all villages. The latter is frequently a problem in the applied market integration work in the WASAT using bivariate correlation coefficients (Harriss).

Conceptually, a seasonal trend in a relatively homogenous geographic region would be the same for all villages in the sample. Divergence from the common trend would either be random or due to other factors (such as market failures), the presence of which we are trying to establish. This provides a means for a practical test of (2) above. This is to estimate seasonal trends separately for the p^*_{it} of each village, in effect $\hat{u}_{it}$, and then to test jointly for equality of seasonal trends across all villages. This involves a joint test of the equivalence of $\hat{u}_{it}$ and $\hat{u}_t$ for all i. Failure to reject the null hypothesis of equality is consistent with the hypothesis of integration of the system of village prices defined as the existence of stable price spreads among locations.

Data

Data are provided by weekly producer price-monitoring surveys carried out from August 1976 to early 1980 by the Nigerian Agricultural Projects Monitoring, Evaluation and Planning Unit (APMEPU), Kaduna, Nigeria. These were part of a much larger evaluation effort by APMEPU of the Funtua and Gusau agricultural development projects in the northern part of the country. The projects cover a continous area of 11,300 km^2 and were implemented with World Bank assistance. They were designed to promote agricultural development through provision of infrastructure, inputs, rural services, and improved seed-fertilizer technologies. Typical farms in the area are two to four hectares. The major food crops, sorghum and millet, accounted for approximately 70% of cropped area in 1977 (Balcet and Candler).

The producer price surveys contain responses by heads of households included in the main agronomic surveys as to the price that either they received for a sale made the day of the interview or that they expected to receive. Although sampling continued through mid-1980, the composition of villages sampled

3

changed in March 1978 in Funtua, and changed with some overlap in March 1977 in Gusau. For the purposes of the present study, a subsample of villages was constructed, covering the period August 1976 to March 1978. Criteria for inclusion in the subsample were set up solely on the basis of availability and continuity of data over the sample period. Only a few villages in Gusau were continuously sampled over the entire period; infrequently sampled villages were unsuitable for present purposes, since the model requires enough observations each season for statistical inference, and thus were dropped. This led to a subsample of 17 villages in Funtua and 5 villages in Gusau. Unfortunately for the study of market integration, survey data on transportation and storage costs were not collected, although grain sales data were obtained on a weekly basis for 118 households surveyed in the 22 villages.

The raw price data contained infrequent but nonnegligible major outliers, perhaps due to keypunching errors or a problem in data collection. The data set was therefore cleaned using a test based on Chebyshev's inequality. Observations that differed from the mean annual observation for the village in question by more than three standard deviations were rejected on the basis that the probability of rejecting a valid observation under this procedure was less than one-ninth. In practice, this led to less than 2% of the data being rejected as unreliable.

The result was a clean, comparable data set of farm-level prices for sorghum and millet covering 22 villages, containing multiple (and differing) numbers of observations for each village and week. Of the 118 heads of household sampled in the 22 villages, 33 were in Gusau. The nonzero observations for each village and week were averaged to get one observation per village per week. Thus, although Gusau was represented by significantly fewer villages, the density of sampling was greater and presumably this led to a better average estimate of producer prices each week.

Definition of Seasons

The first area of investigation concerns the delimitation of seasons. In northern Nigeria, two possible classifications of seasons come to mind for present purposes. One is the separation of the year into a wet season (June to September) and a dry season (October to May). Agricultural production is concentrated in the former; frequently road transport to outlying areas is difficult at that time. An alternative classification is the one adopted by Hays, based on grain sales behavior. This divides the year into two six-month periods: harvest (August to January) and postharvest seasons. Harvesting of food crops begins with early millet in August, includes sorghum in October and ends with cowpeas in late November to early December. Cotton is harvested thereafter. Hays found that in 1970/71, small farmers accomplished 50% of their millet and 69% of their total sorghum sales in the harvest season, while the corresponding figure for large farmers was 30% for both crops.

The present data set of 118 households, which may include a disproportionate number of larger farmers, shows that from August 1976 to July 1977, 48% of millet and sorghum sales occurred in the harvest season, from August to January. Similarly, during calendar 1977, 43% of millet and sorghum sales occurred during the June to September rainy season. Clearly, substantial grain sales occur in Northern Nigeria in both seasons. Available data do not support the conclusion that the volume of grain sales is substantially less during the more restrictively defined rainy season. Peak sales for the combined sample occurred in September and December in both 1976 and 1977 (the harvest season) and also in May in 1977, in the middle of the postharvest period.

Even if sales volumes are not very different by season, it is possible that market actors are different. Several authors have suggested that grain transactions in rural areas of Northern Nigeria in the harvest period occur to a greater extent among farmers than is the case in the postharvest season, when long-distance traders and middlemen play a relatively more important role (Scott, Hays). This seasonal difference appears to be a function more of prevailing grain prices than of climatic factors, although further research is warranted here.

The behavior of nominal grain prices provides another clue to the correct definition of seasons for present purposes. Plots of the raw price data for the twenty-two villages shows that prices tended to rise steadily during the postharvest season, but were highly variable in the harvest season. This is characteristic of patterns portrayed for earlier years by Scott and Hays.

Finally, the data available do not permit direct verification of seasonal differences in transport and transactions costs. While it seems reasonable that these may be higher in the rainy season because of transportation difficulties, no village in the sample was more than 10 kilometers from a rainy season road. A survey of the literature quoting grain transportation costs in Northern Nigeria could find no differentiation of costs on a seasonal basis. The assumption adopted, therefore, is that differences in transport and transactions costs are a function of the same variables that determine the choice of seasons: sales volume, grain price variability, and type of market actors.

Therefore, the price data set was subdivided into three seasons, corresponding to the major dates of the agricultural calendar from August 1976 to mid-January, 1978. The two harvest seasons were defined as running from early August to mid-January. The intervening postharvest season covered mid-January to early August 1977. The period covered during the 1978 postharvest season was too short to be usable. The unknown variance of the z terms in (1) above is assumed to be constant within seasons for each crop and village but not necessarily across seasons, crops, or villages. Furthermore, any variation in price spreads between any pair of markets within a given season, other than random variation, is assumed to be caused by factors other than variation in transportation and transaction costs.

The Tests

Given these assumptions, the model in (4) above was explored by removing village-level seasonal trends, the u_{it}, by a separate ordinary least squares (OLS) polynomial regression of p_{it}^* on powers of t for each crop, village, and season. Initially, a high order of polynomial was used, since including too many terms leads only to inefficiency based on a low number of degrees of freedom. However, using too low a polynomial introduces serial correlation into the residuals. The degree of polynomial for each village was reduced until the t-statistic of the last term became significant at the 0.10 level. In practice, the 1976/77 postharvest season could be acceptably modeled with a quadratic, while both harvest seasons needed to be modeled with a sixth-order polynomial. To avoid introducing a differential element between seasons in data to be used to test for seasonal differences, all residuals used for the tests below were generated with a sixth order polynomial, regardless of season. The high order of polynomial required for the harvest season stems from a bimodal price structure observable in the six months following the early millet harvest; the reasons for this structure will be explored later. The result is a set of residuals generated separately for each village: $\hat{z}_{it_{ors}}$.

The test of market integration in the system is then to test for equality of trends across villages, but this requires an estimate of the common seasonal trend for each crop and season in order to retrieve estimates of the $\hat{z}$ residuals from a single joint regression. However, homogeneity of variance of food grain prices across villages is a riskier assumption than constancy within a given village. Thus, it was necessary to test explicitly this proposition. This was done using a derivative of Bartlett's test. The appropriate statistic, asymptotically distributed chi-square $(m - 1)$ is

$$(5)\quad M = \frac{(T - m)\ln \sigma^2 - \sum_{i=1}^{m} (T_i - k_i) \ln \hat{\sigma}_i^2}{1 + [1/3(m - 1)]\left[\sum_{i=1}^{m} 1/(T_i - k_i) - 1/(T - m)\right]},$$

where m is the number of villages, σ^2 is the estimated variance of the pooled OLS regression residuals, T_i is the number of observations in the ith sample and the total number when the subscript is dropped, and k_i is the number of regression parameters estimated from the ith sample (village).[1]

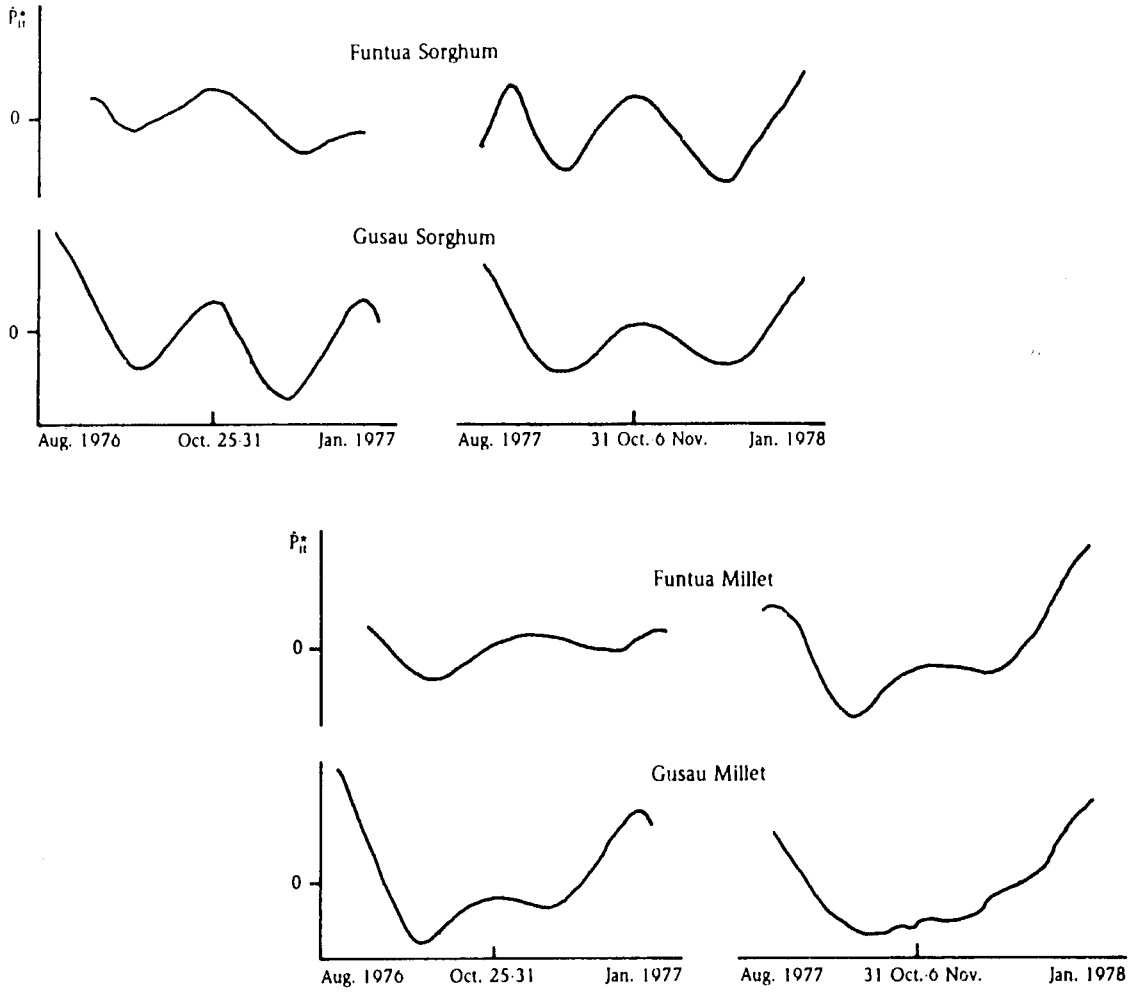
The tests showed that it was possible to reject the null hypothesis of homogeneity of variance across villages for a given crop and

[1] This formulation differs from that given in Judge et al. by inclusion of the k_i, a weighting term that adjusts the degrees of freedom for the fact that regression residuals were used. Asymptotically, this adjustment makes no difference. In small samples, it will increase both the numerator and the denominator with the overall effect on M dependent on specific cases. In practice, if the null hypothesis of equality of variances is rejected (M is large), it is unlikely to make much difference.

season at one percent in many cases. Thus, it was apparent both that prices had different distributions in different villages and that the heteroscedasticity resulting from simple pooling of data across villages could lead to an overestimate of the variance of the pooled error term, which would in turn vitiate tests based on this variance. It was also apparent from similar tests that detrended food grain price variances were significantly more homogenous within either the Funtua or Gusau project regions than across the whole sample. Accordingly, pooled polynomial regressions were run for each season for villages in each project separately, using standard Aitken weighting matrices composed of the reciprocals of the standard deviations of the OLS residuals from separate regressions for each village and applying generalized least squares (GLS) to the villages pooled by project. Plots of the estimated prices by project with the seasonal means removed are contained in figure 1.

The main hypothesis of the experiment can now be directly tested: to wit, that seasonal food grain price trends are the same across villages. This is done separately for Funtua and Gusau and for each crop and season by the statistic:

$$(6)\quad f = \frac{\left(\sum_i \sum_t \hat{z}^2_{it_{OLS}} - \sum_i \sum_t \hat{z}^2_{it_{GLS}}\right)/r}{\sum_i \sum_t \hat{z}^2_{it_{GLS}} \Big/ \sum_i (n_i - k_i)},$$



Note: $\hat{P}^*_{it} = \sum_{k=0}^{6} \beta_{k_{GLS}} t^k$

Figure 1. Plots of estimated prices during harvest periods with seasonal means removed by village

6

Table 1. Seasonal Differences in the Homogeneity of Price Trends across Markets, by Crop and Project

Season	Crop	Project	F	Degrees of Freedom	H_0[a]
Postharvest 1976/77	Sorghum	Funtua	0.35	(112,131)	Fail to reject
		Gusau	1.86	(28,86)	Fail to reject
	Millet	Funtua	0.27	(112,12)	Fail to reject
		Gusau	1.41	(28,73)	Fail to reject
Harvest 1976	Sorghum	Funtua	2.36	(112,244)	Reject
		Gusau	4.11	(28,73)	Reject
	Millet	Funtua	2.46	(112,214)	Reject
		Gusau	2.52	(28,77)	Reject
Harvest 1977	Sorghum	Funtua	1.29	(112,55)	Fail to reject[b]
		Gusau	5.44	(28,82)	Reject
	Millet	Funtua	2.34	(112,54)	Reject
		Gusau	10.74	(28,79)	Reject

[a] H_0: $b_{ik} = b_k$ for all markets i in the subsample and parameters k, tests at 1% significance level.
[b] The critical value for rejection is circa 1.35 at 10%.

which is distributed $F(r, \Sigma_i(n_i - k_i))$ under the null hypothesis of equality of trend (Ho), where r is the number of independent parameter restrictions applied to the pooled regression [i.e., $7 \times (17 - 1)$ for Funtua], n_i is the number of observations in village i and k_i is the number of parameters estimated for village i (generally 7).[2]

This test statistic is valid under Ho, but conceptual problems arise if Ho is rejected. The latter implies that dissimilar trends were pooled, leading to serial correlation of the GLS residuals, implying that the variance of the GLS residuals used in the numerator is biased.

However, the type of serial correlation introduced by pooling divergent village trends into a common trend is almost certainly positive for each village in the pooled regression, implying that it is positive for the pooled regression as a whole. This implies that the estimated variance of GLS residuals is biased downwards (Judge et al.). Because the OLS residuals are estimated separately by village and are therefore not affected by this problem, it follows that the net effect of pooling dissimilar trends in (6) above is to bias the test towards accepting Ho. Thus, rejections of the null hypothesis of equality of trend, and thus of market integration, are not affected. In fact, the problem is akin to the well-known Behrens-Fisher problem of testing for the equality of two means with unequal variances. Results are asymptotically efficient, with a bias towards accepting the null hypothesis in small samples (Kendall and Stuart, pp. 153–65).

Results

The results of the tests are shown in table 1. They show a striking dichotomy between the 1976/77 postharvest season and the two harvest seasons. The null hypothesis of equal seasonal trends cannot be rejected for any crop or project in the postharvest season. A common seasonal trend with constant seasonal price spreads among villages cannot be ruled out. However, the reverse is true of the harvest season data. In seven out of eight cases, the

[2] The combined use of the pooled GLS and individual OLS residuals in this statistic does not present a problem, since weighting an individual OLS run by the constant GLS weights for the village in question will lead to the same sum of squared residuals as simple OLS on the individual village data.

null hypothesis of equality of trends can be rejected at one percent. The eighth case, the 1977 harvest season for sorghum in Funtua, only narrowly fails at 10%.

Given that the possible downward bias in the estimated variance of the pooled GLS residuals only strengthens the results in table 1, rejection of the hypothesis of equal seasonal trends across villages is straightforward for both harvest seasons. Given the formulation and assumptions of the model, it would appear that the system of food grain markets is not well integrated. On the other hand, acceptance of the null hypothesis—that the system is integrated—is possible for the postharvest season, subject to the statistical caveats expressed above.

Of course, if price variation in the postharvest season was especially high, it would help explain why it is more difficult to reject the null hypothesis of equal seasonal trends. Yet results (not shown) demonstrate conflicting evidence for Funtua and Gusau in this regard, despite uniform results for market integration in table 1.

The pooled $\hat{z}$ residuals by crop and project can be used to show inequality of seasonal variances. The procedure in equation (3) eliminates constant effects attributable to the mean price level and market-specific deviations from that mean. Therefore, the remaining variation in equation (4) is either due to variation across time periods attributable to the common seasonal trend u_t, or variation around that trend over time and markets, z_{it}. Therefore, variances computed from the residuals estimated within each crop and project by the GLS regressions pooled across markets also provide an estimate of the variance of food grain prices around the common seasonal trend of the markets concerned. These estimated variances are distributed chi-square and can be compared pair-wise by F-tests when divided by the appropriate degrees of freedom. Whereas the postharvest season variance for sorghum prices in Funtua is larger than each of the two harvest seasons at the 5% significance level, this is true of only one harvest season for millet and was not true at all in Gusau.

In sum, the weekly farm-level price data for August 1976 to January 1978 from Funtua and Gusau suggest three conclusions with respect to the price integration of local food grain market systems in northern Nigeria. First, the differences in price behavior between the adjacent projects of Funtua and Gusau were sufficiently great so as to preclude pooling them jointly for consideration as a unified market system. Second, within each project, there are apparent seasonal differences in food grain market integration between the harvest and postharvest periods. Market integration is defined here as the existence of stable price spreads among markets in a given season, despite considerable variation in prices. Third, markets are clearly not well integrated in Funtua or Gusau in the harvest period. The reverse conclusion is not necessarily valid for the postharvest period, but at least the evidence is consistent with the hypothesis of market integration thus defined.

A fourth result is the finding that the variance of detrended and deseasonalized food grain prices is different in the harvest and postharvest seasons. One immediate consequence is to suggest caution when using bivariate correlation coefficients as a statistical technique to investigate market integration. Even when used with detrended and deseasonalized data, their statistical interpretation is based on the assumption that variances are constant for each series. Thus, even a "low" value cannot be used as an argument to show the absence of integration, since under these circumstances there is no criterion of what is low.

A fifth, probably related, result of the experiment is the unexpected finding of a general bimodal price trend for both millet and sorghum during both harvest seasons, as shown in figure 1 above. A priori, this is contrary to economic theory: if it is a generally known and expected trend and if credit and storage facilities work in any manner at all, it is hard to understand why anyone would buy at the bump during the second half of October, or sell in the troughs on either side. Since the sales data show that grain sales do occur throughout this period, it is reasonable to suppose that the existence of the bump is somehow linked to the finding of lesser market integration during the harvest season. This would most likely occur through some third force affecting both phenomena jointly. A better understanding of the reasons for the bump would lead to a better understanding of how markets function.

One possibility is that the October price bump is specific to Funtua and Gusau in 1976–78, considerably weakening its analytical interest. If this phenomenon recurs frequently,

it is reasonable to infer that a relatively lower degree of market integration in the harvest season also recurs in other years and places, although a more conclusive demonstration will have to await further research based on new primary data. Nevertheless, published secondary data can be used to trace food grain price behavior in October on a wider scale.

Generality of the October Price Bump and Possible Explanations

One possible explanation for both the relatively lower market integration and the October-centered bump in the 1976 and 1977 APMEPU harvest season data is that an important Islamic festival (*id-al-Kabir*) occurred at that time. This is a period typically associated with increased food expenditures, where economy may be secondary to last minute necessity in consumer minds. Yet, secondary price data suggest that this is not the explanation.

Available information consists of 11 years of monthly wholesale prices for millet and sorghum in 15 towns in Northern Nigeria, published by Hays. The data covers the periods 1958 to 1965 and 1969 to 1971 and come from Nigerian crop and weather reports. Approximately one-third of the 330 yearly price series exhibit the October-centered bump of figure 1 in some form. The major Islamic festivals are based on the lunar calendar, and thus occur every year a dozen days earlier on the Gregorian calendar. *Id-al-Kabir*, for example, occurred in mid-March in 1962. Yet the October bump can be found in Hays' data in this and other years where no festivals were held in October. Thus the bump does appear to be associated with the harvest season.

The bump can be explained in a number of ways. One explanation might be distress sales in the first trough following the early millet harvest, followed by withdrawal once immediate cash needs are met. A more likely explanation is that two kinds of grain, old and new, are being sold during the harvest period, without being differentiated by the price series but being differentiated by the consumer. When the condition of the harvest is clear, farmers rush to sell off the old grain in store, which may lose value when the new crop is in. Thus, the first trough possibly measures the price-depressing effects of the early millet harvest compounded by the selling off of old sorghum; the bump occurs as more valuable new grain replaces old grain in the market and the second trough occurs as the full new harvest takes effect. The mix of new and old grain in each village is probably quite different, leading to apparent variation in geographic price spreads, when in fact different commodity bundles are involved. If correct, this would by itself explain why price correlations for food grains in the WASAT are typically so poor despite widespread belief that systems are competitive. Under this scenario, varying mixes of apples and oranges are being correlated with each other. The problem would then exist for any analysis based on misleading price data, the present one not excepted, and not with the markets.

Whatever the explanation, it is clear that analysts need to focus on those factors in food grain marketing that vary between harvest and postharvest seasons, and particularly need to explain marketing behavior in the harvest season. Variables that may be relevant in this regard concern differences in (*a*) volume of sales and purchases; (*b*) the frequency, size, and variability of market presentations; (*c*) market agents and the destination of grain traded; and (*d*) characteristics of food grains sold.

One of the most plausible explanations of seasonal differences in market performance is one suggested by Scott and Hays more than a decade ago and still largely unstudied in a market efficiency context. This is the hypothesis that grain transactions in rural areas in the harvest period occur to a greater extent among farmers than is the case in the postharvest season. If true, the implication may be that in the postharvest season traders play a relatively greater role in transmitting price signals among villages.

Summary and Conclusions

The variance components method permits statistical inference from a sample of time series of market prices concerning seasonal and regional differences in the variance of food grain prices. Evidence of the former in northern Nigeria casts even more doubt on the usefulness of the bivariate correlation coefficient when used across seasons to make inferences about integration of pairs of markets. Furthermore, evidence of differences in the variance of food grain prices between two contiguous agricultural project areas that together cover only 11,000 square kilometers suggests that

9

the land area associated with a given system of markets may be relatively small in northern Nigeria, even though influences among market systems are presumably important.

The strong evidence for seasonal differences in market integration, defined as the existence of stable price spreads between markets despite temporal price changes within seasons, suggests the need to look carefully at marketing behavior that differs across seasons. Research should focus on explaining the distinctive bump in prices that occurs in the middle of the harvest season in some years, since it appears likely that this is associated with the finding of lower market integration at that time.

[*Received February 1984; final revision received September 1985.*]

References

Balcet, J. C., and W. Candler. *Farm Technology Adoption in Northern Nigeria*, vol. 2. Mimeographed. Washington DC: World Bank, West Africa Projects Dep., Feb. 1982.

Balestra, P., and M. Nerlove. "Pooling Cross-Section and Time-Series Data in the Estimation of a Dynamic Model: The Demand for Natural Gas." *Econometrica* 34(1966):585–611.

Delgado, C. L. "Foodgrain Market Integration in Northern Nigeria: Implications for Subregional Equity Concerns." Paper presented at the International Workshop on Agricultural Markets in the Semi-Arid Tropics, International Center for Research in the Semi-Arid Tropics, Patancheru, India, 24–28 Oct. 1983.

Ejiga, N. O. "Economic Analysis of Storage Distribution and Consumption of Cowpeas in Northern Nigeria." Ph.D. thesis, Cornell University, 1977.

Granger, C. W. J. "Investigating Causal Relationship by Econometric Models and Cross-Spectral Models." *Econometrica* 37(1969):424–38.

Gupta, S., and R. Mueller. "Analyzing the Pricing Efficiency in Spatial Markets: Concept and Application." *Eur. Rev. Agr. Econ.* 9(1982):301–12.

Harriss B. *The Marketing of Foodgrains in the West African Sudano-Sahelian States: An Interpretive Review of the Literature*. Hyderabad: ICRISAT, Econ. Program Progress Rep. No. 31, Jan. 1982.

Haugh, L. D. "Checking the Independence of Two Covariance Stationary Time Series: A Univariate Residual Cross-Correlation Approach." *J. Amer. Statist. Assoc.* 71(1976):378–85.

Hays, H. M., Jr. *The Marketing and Storage of Food Grains in Northern Nigeria*. Samaru, Zaria: Institute of Agr. Res. Misc. Pap. No. 50, 1977.

Hays, H. M., and J. H. McCoy. "Food Grain Marketing in Northern Nigeria: Spatial and Temporal Performance." *J. Develop. Stud.* 14(1977):182–92.

Jones, W. O. "The Structure of Staple Food Marketing in Nigeria as Revealed by Price Analysis." *Food Res. Inst. Stud. in Agricultural Economics, Trade and Development*, no. 2(1968), pp. 95–123.

Judge, G. G., et al. *The Theory and Practice of Econometrics*. New York: John Wiley & Sons, 1980.

Kendall, Maurice, and A. Stuart. *The Advanced Theory of Statistics*, 4th ed., vol. 2. New York: Macmillan Co., 1979.

Scott, E. P. "Farmers' Participation in the Marketing of Food Crops in Two Rural Hausa Markets." *Savanna: A journal of the environmental and social sciences published at Ahmadu Bello University, Zaria* 7(1978):107–26.