Testing the link between public intervention and food price variability: evidence from rice markets in the Philippines

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ABSTRACT. Monthly price and stock data are used to test the influence of Philippine government buffer-stock programs on seasonal and annual variability of producer and consumer rice prices. The period examined is 1974-1990. NFA stock changes are shown to have had some stabilizing influence on seasonal and annual price changes, but the magnitude is small and not statistically significant.

Keywords: Philippines, food price policy, government buffer stock programs, rice

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1. Introduction

The goal of reducing poverty and hunger has made food price policy a high priority in most developing countries (Claessens and Duncan, 1993; Rao, 1989; Sahn, 1989). Many governments pursue policies aimed at narrowing the band between consumer and producer prices and at stabilizing prices within and between years (Timmer, 1989; Islam and Thomas, 1996). Attempts to influence food prices have been particularly pervasive in Asia, where the dominant food is rice. Barely five percent of the world's total rice production is traded (Berck and Bigman, 1993). This "thinness" is a primary cause of price instability in the world rice market, and a frequent justification for government intervention in rice markets (Timmer and Falcon, 1975).

In the Philippines, where rice is by far the most important crop in terms of food use and production value, the Philippine National Food Administration (NFA) has sought to establish stocks and control imports in an effort to stabilize domestic prices during the past two decades.[1] It does so at large economic cost (Unnevehr, 1983). Generally, the NFA intervenes at two distinct times of the year. At the peak of the wet-season harvest, when farmgate prices are falling, the NFA buys rice from producers at a set price. During the dry season, when market supplies are low and prices are high, the NFA releases stocks to licensed outlets to be resold at established maximum retail prices.

The NFA's stated goals have been to ensure sufficient rice stock for supply shortfalls, to protect rice farmers' incomes, and to make rice affordable to urban consumers (PhilRice-BAS, 1995). However, observers have noted that Philippine policies have tended to raise rice prices above their free-trade level (STAT-USA, 1996), and have been less successful in stabilizing

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[1] As Corpuz (1997) points out, pricing policies in the Philippines have a long history. Imports have been modest, however, and have rarely exceeded five percent of consumption, even during periods of low domestic production (STAT-USA, 1996).

prices than the policies of other countries in the region. For example, Dawe and Timmer (1991) compare the NFA's actions to those of Indonesia's BULOG and argue that the NFA has had comparatively less success influencing rice prices, in part because it announced both target prices and target procurements at irregular times during the year.[2]

In this paper we use producer and consumer price data to test whether government buffer stock actions stabilized rice prices in the Philippines during the period 1974-1990. We test the impact of stock changes by econometrically estimating the magnitude of seasonal and annual price variation, with and without explicit account of NFA stock changes. We find a very small and statistically insignificant stabilizing effect. Our results demonstrate that NFA interventions have been largely unsuccessful at driving prices, and have not substantially reduced price variability.

2. Model

To assess how government interventions might affect prices, consider a country that neither accumulates government stocks, nor engages in international trade. A simple model of competitive stock holding in the market, following Williams and Wright (1991), would predict price and stock level patterns illustrated by the solid lines in Figure 1. At the conclusion of the harvest, supply reaches its peak and prices fall to their lowest level. At the start of the subsequent production period, price begins to rise.[3] During the period between harvests, price reaches its highest level, as stocks fall to their lowest level. Private stock releases during the

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[2] By generating greater price uncertainty for traders, such actions tend to raise the cost of private investment in stocks (Shively, 1998). As a result, government policies have the potential to increase price instability by undermining private stockholding.

cycle vary according to current prices, expected future prices, and storage costs. At the end of one harvest a grain trader may decide to store all or part of his grain in expectation that the price will rise again prior to the subsequent harvest. This prevents the minimum, post-harvest price in one season from falling below the post-harvest price in the next season. As price peaks, traders perceive an incipient price decline following the next harvest, and private stocks approach zero.

Competitively determined imports or stock releases during the period preceding a harvest should tend to reduce seasonal price rises. Provided the cost of importing grain is greater than the lowest domestic price observed before the first harvest, free trade at a constant world price would lead to the price and stock level trends illustrated by the dashed lines in Figure 1. The dotted boxes illustrate periods without imports. The price line shows that private storage would be undertaken only until the period of imports began. Stock releases during the importing period would not affect price but would instead substitute for commercial imports.[4]

Our goal is to investigate whether NFA interventions stabilized prices over the sample period. We begin by asking whether the NFA successfully stabilized seasonal price fluctuations, i.e. whether NFA actions reduced average month-to-month price changes. To investigate this issue, we estimate regressions for farmgate and retail prices based on the price equation:

$$\Delta P_t = \beta_0 T + \beta_1 NFA_t + \sum_i \alpha_i M_i + \varepsilon_t \qquad (1)$$

where ΔP_t is the monthly change in price (either farmgate or retail), T is a unit-step time trend, NFA_t is the monthly change in government stock, and M_i is a monthly dummy variable to

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[3] The rate of increase is determined by storage costs and the rate of stock depletion. The second harvest, were it unexpectedly large, could temporarily depress price, but price would subsequently rise due to costs of storage.

[4] The stock release necessary to achieve the target price will depend on the price elasticity of demand. For the Philippines, Bouis (1982) estimated a long-run price elasticity of demand for rice of 0.5.

account for systematic seasonal changes in prices. In accordance with the logic outlined above, if NFA purchases raised prices on average (and stock releases lowered them), one would expect the sample estimate of β_1 to be positive and statistically significant. More importantly, if NFA actions were successful in reducing seasonal variability, the α_i estimates on each month's price change would be larger in a regression that includes the *NFA* stock change than in one that does not control for NFA actions.

Investigations based on equation (1) test the relationship between seasonal price changes and NFA activities each month throughout the year. We also consider the possibility that NFA actions reduced the sensitivity of prices to harvest size, and thereby stabilized inter-annual price fluctuations. To examine this question, we estimate annual regressions of the form:

$$\mathrm{d}P_t = \beta_1\, \mathrm{d}Q_t + \beta_2\, \mathrm{d}NFA_t + \epsilon_t \qquad (2)$$

where $\mathrm{d}P_t$ is the deviation of the period's price from the sample average (July-Aug. in the case of retail prices and Sept.-Oct. in the case of farmgate prices), $\mathrm{d}Q_t$ is the deviation of the critical period's harvest from the sample average (Jan.-June in the case of retail prices and July-Dec. in the case of farmgate prices), and $\mathrm{d}NFA_t$ is the NFA stock change in the period of interest. As with the seasonal stabilization test in equation (1), the sign and significance of the estimated coefficient for β_2 serves as a test of whether NFA stockholding is itself responsive to price. But more importantly, we can use equation (2) to test the link between price sensitivity to harvest size and NFA actions. If NFA stock changes successfully offset production fluctuations, the estimated value of β_1 will be larger in a regression that includes the d*NFA* variable than in one that does not account for government stock change.

3. Data

Data on national, i.e. regionally-weighted aggregate prices and NFA activities over the period 1974-1992 come from the Philippine Bureau of Agricultural Statistics (PhilRice-BAS, 1994). Prices were deflated using CPIs reported in Boyce (1993) and the UN Statistical Yearbook (1994). Monthly NFA stock-level changes are used in place of actual NFA purchases and releases, because the latter could not be reliably identified. Data on production come from USDA (1996).

4. Results

Regression results for models of monthly price changes based on equation (1) are presented in Table 1 for the period 1974-90. Columns 1 and 2 of the table contain results for regressions of farmgate prices; columns 3 and 4 contain results for regressions of retail prices.[5] The time trend coefficients indicate that price changes did not change systematically over the sample period. A strong seasonal pattern in price changes is apparent in both farmgate and retail prices. In the case of farmgate prices, monthly price changes were negatively correlated with NFA stock changes at a statistically significant level over the sample period. In other words, results confirm that the NFA intervened by purchasing stock at times when farm prices were low. In the case of retail prices, monthly price changes were not correlated with NFA stock changes at standard significance levels.

The seasonal price patterns implied by these regressions are illustrated in Figure 2, where for each regression, the values of the monthly dummy variables have been plotted. The solid

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[5] The use of first differences in prices (i.e. p_t-p_{t-1}) as the dependent variable is justified on the basis of stationarity tests. Using Dickey-Fuller tests, the null hypothesis of a unit root could be rejected for both farmgate or retail price changes. Test statistics were -3.76 and –4.45, respectively, for models with constant and trend (vs. a 10 % critical value of –3.13).

lines in the figure represent monthly changes computed in regressions that did not control for NFA interventions. The dashed curves in the figure represent monthly changes computed in regressions that explicitly include changes in NFA stock. Hence, the solid lines show price changes achieved with NFA intervention. The main finding is that NFA interventions did have the desired effect on monthly price changes during the study period, but that the magnitude of the effect was very small. For example, when we compare models with and without NFA actions, we find that with the NFA the average monthly rise in the retail price is only about 10% smaller between February and July—the pre-harvest season. Similarly, with the NFA the average seasonal drop in the farm gate price is only about 11% smaller between September and November—the harvest season. Implied differences in cumulative price changes are correspondingly small. Furthermore, these differences are not statistically significant, in the sense that we cannot reject a hypothesis that the parameter values differ individually or collectively between the regressions.[6]

Table 2 contains results from tests of inter-annual stabilization, conducted using two sets of regressions based on equation (2). These regressions were designed to investigate the possibility that NFA actions were successful in reducing the sensitivity of post-harvest farmgate prices and pre-harvest retail prices to the size of the harvest. The first pair of regressions in

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[6] In separate regressions not reported here, we allowed the impact of the NFA stock change on price changes to vary for each month by including 12 separate month × NFA stock change variables in the regressions (in addition to the monthly dummy variables). Although few of the individual slope estimates were individually significant at standard test levels, the hypothesis of joint significance could be accepted for both the farmgate and retail price regressions. Formally, tests of the hypotheses that $[\boldsymbol{\gamma}] = [\mathbf{0}]$ in the regression:

$$\Delta P = \alpha \text{ time } + \beta' \text{Month } + \gamma'(\text{Month*}\Delta\text{NFA}) + \epsilon$$

gave F-statistics of 1.09 and 0.84 for the farmgate and retail price regressions, respectively. The statistics are distributed $\chi^2(10)$, and in each case fall well below the 95% critical value of 18.3. This suggests that impacts of NFA interventions were likely not uniform across the year.

Table 2 focuses on the relationship between changes in the farmgate price in the period immediately following the wet season harvest and production and NFA stock changes. The second set focuses on the relationship between changes in the retail price during the pre-harvest season and production and NFA stock changes in the preceding period.

The dependent variable in the farmgate regression is computed as the average September-October farmgate price for the year, expressed as a deviation from the sample average for all September-October prices. The regression indicates that farmgate price changes were negatively correlated with production changes over the study period. In elasticity terms, a one-percent increase in the deviation of the annual harvest from its sample average was associated with a 0.10% reduction in the post-harvest farmgate price change. However, the NFA stock change variable is not significantly correlated with price changes.[7] As before, adding the NFA stock change variable to the regression has a small impact on the relationship between production and prices in the harvest period. NFA interventions appear to lower the sensitivity of price to harvest size by a similar magnitude as before (on the order of 10%), but the differences are not statistically significant and therefore fail to support a hypothesis that post-harvest NFA stock purchases were sufficient to stabilize prices inter-annually.

5. Discussion and Conclusions

This paper examined the relationship between government buffer-stock programs and the level and variability of rice prices in the Philippines. Producer and consumer prices were used to estimate two econometric models to test the relationship between price changes and government

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[7] Regressions on annual data in which NFA stock changes in pre-harvest and post-harvest periods were regressed on deviations from average production suggest that, on average, the NFA accumulated larger stocks in years with above-average production. The results reported in tables 1 and 2 suggest these actions were insufficient to alter seasonal price patterns.

stock actions over the period 1974-1990. Significant correlation was observed between NFA stock changes and price changes, and stock changes did provide some seasonal and inter-annual stabilization, but the effect was very small and was not statistically significant. Hence, arguments that NFA actions altered the seasonal pattern of prices or its inter-annual variability are valid but not compelling. These results raise questions regarding the overall effectiveness of NFA activities, especially with regard to producer prices. Given the high budgetary cost of NFA activity, future NFA actions should be weighed against alternatives use of resources that might prove more effective in protecting consumer and farmer incomes.

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Table 1. Regressions of farmgate and retail price changes

	Farmgate		*Retail*	
Time	0.24e-04 (0.42e-04)	0.16e-04 (0.40e-04)	0.40e-04 (0.73e-04)	0.31e-04 (0.72e-04)
January	-0.0916 (0.0102)	-0.0833 (0.0106)	-0.2252 (0.0188)	-0.2155 (0.0199)
February	0.0257 (0.0098)	0.0292 (0.0098)	0.0141 (0.0182)	0.0183 (0.0183)
March	0.0069 (0.0098)	0.0076 (0.0097)	0.0252 (0.0182)	0.0032 (0.0181)
April	0.0083 (0.0099)	0.0097 (0.0097)	0.0102 (0.0182)	0.0118 (0.0182)
May	0.0145 (0.0099)	0.0165 (0.0098)	0.0172 (0.0183)	0.0196 (0.0183)
June	0.0194 (0.0099)	0.0217 (0.0098)	0.0300 (0.0183)	0.0326 (0.0183)
July	0.0208 (0.0099)	0.0233 (0.0098)	0.0450 (0.0183)	0.0479 (0.0184)
August	0.0005 (0.0099)	-0.0004 (0.0098)	0.0302 (0.0184)	0.0290 (0.0183)
September	-0.0356 (0.0100)	-0.0391 (0.0099)	0.0198 (0.0184)	0.0157 (0.0185)
October	-0.0430 (0.0100)	-0.0495 (0.0102)	-0.0292 (0.0184)	-0.0366 (0.0191)
November	-0.0022 (0.0100)	-0.0806 (0.0099)	-0.0211 (0.0185)	-0.0194 (0.0184)
December	0.0186 (0.0100)	0.0272 (0.0106)	0.0060 (0.0185)	0.0160 (0.0197)
NFA stock change	–	-0.0387 (0.0171)	–	-0.0450 (0.0315)
ρ	0.0177 (0.0682)	-0.0104 (0.0682)	-0.0552 (0.0681)	-0.0712 (0.0680)
R^2	0.45	0.47	0.49	0.50
number of observations	215	215	215	215

Notes: Standard errors in parentheses. Reported R^2 is calculated from raw sample moments due to suppression of intercept. Dependent variable is first difference in prices. The hypothesis that estimated coefficients on monthly dummies in “with NFA” regressions are the same as monthly dummies in “without NFA” regressions cannot be rejected in either individually paired or joint tests in both farmgate and retail regressions.

Table 2. Regressions of seasonal price changes on production and NFA activity

	Farmgate price change		*Retail price change*	
Change in production (metric tons)	-0.1304† (0.0491)	-0.1137† (0.0546)	-0.16e-06†† (0.79e-07)	-0.15e-06†† (0.07e-07)
Change in NFA stock (metric tons)	–	-0.0887 (0.1196)	–	0.8926 (0.3152)
R^2	0.29	0.32	0.19	0.46
number of observations	18	18	18	18

Notes: Standard errors in parentheses. Reported R^2 is calculated from raw sample moments due to suppression of intercept. Price and production variables are expressed as deviations from sample means. The symbols † and †† indicate the coefficients are not statistically different at a 95% confidence level.

Figure 1. Seasonal price-stock relationship with and without intervention

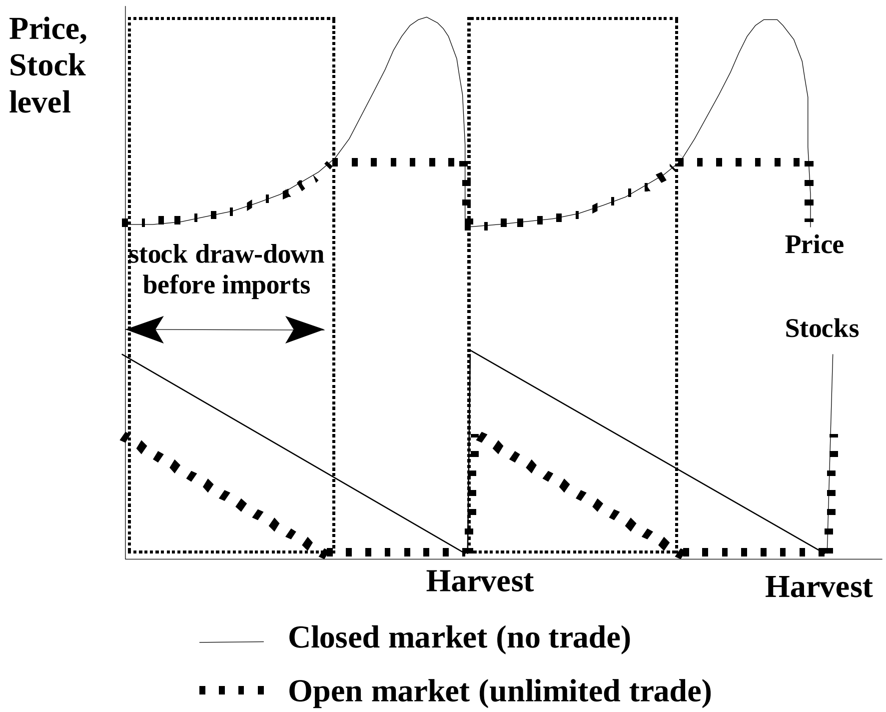

Figure 2. Pattern of monthly price changes based on regression estimates