International and Internal Market Integration in Indian agriculture: A study of the Indian Rice Market

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Abstract

There has been concern about the effectiveness of India's agricultural policy reforms adopted in recent years as part of the overall policy liberalisation process. These concerns have been strengthened by studies of spatial market integration of major agricultural commodity markets, such as the rice market, which have concluded that Indian agricultural markets remain largely segmented and fragmented. These sudies, however, have ignored possible structural breaks due to reform policies adopted since the early 1990s and the possible impact of world markets on domestic price movements. We show that the major reforms of the Indian rice market in 1994 has had a major impact on market integration, leading to much faster price convergence between domestic and international prices. The pace of price convergence is influenced by quality of infrastructure in the states and whether they produce market surpluses, possibly because of the asymmetric nature of foreign trade liberalisation in rice.

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1. Introduction

After decades of pervasive restrictions on both international and internal trade, India started to implement some limited policy reforms in the 1980s, and sharply accelerated the reform process in the early 1990s. Major trade policy liberalisation and macroeconomic reforms have transformed the economic and policy environment and the country has clearly embraced economic globalisation. The results have been quite dramatic: for well over a decade now, India has been experiencing unprecedented rates of overall economic growth.

But there have been widespread concerns about impact of these policy liberalisation measures on the agricultural sector and the rural economy, and the consequences for poverty and food security. While the reform process has certainly impacted on agriculture, there have been concerns about the pace, scope and effectiveness of agricultural sector reforms. These have been given weight by recent studies (see, for example, Jha et al, 2005) that have argued that reforms have not been effective in addressing the segmentation of domestic markets, which hinders the emergence of competitive market structures, and insulate them from each other as well as from international markets, constraining the achievement of improved market efficiency.

These conclusions are both surprising and disturbing. If correct, they would imply that, despite the many major regulatory reforms announced by the government, no significant change has occurred in the way the major agricultural markets function in India. The manner and extent of price movements among the various domestic markets, and between domestic markets and international markets is an important indicator of the effectiveness of the reform measures.

If government interventions distort price signals in spatially separated internal markets, domestic prices may not converge efficiently. Market segmentation is also consistent with non-competitive markets and trade liberalisation at the border does not have expected impacts because international price changes are not efficiently transmitted to domestic markets.

In this paper we report results of the first stage of an analysis of the evolution of internal and international market integration of a major cereal grain market in India, the rice market, since 1980.[1] The overall objective of this study is not only to assess the level of integration, but to see how it has changed over time, and to examine the contribution of central and state government policy reforms.

2. Agriculture Sector Reforms and Domestic Trade in Rice

The focus of the Indian reform process was initially on the manufacturing sector, but gradually extended to the other sectors including agriculture. A key component of reforms have been measures to lower restrictions on the internal movement of agricultural commodities and the liberalisation of foreign trade. The pervasive restrictions that have inhibited free movement of key agricultural products (particularly cereal grains) across various administrative regions, are well known and extensively documented.[2] The reform measures have included changes to the Essential Commodities Act (1955) which regulated internal trade in major agricultural products (e.g., removal of the licensing requirements and stocking limits for the wholesale and retail trade), and abolition of selective credit controls used to regulate institutional credit to traders. As a result of reforms, state trading activities, once the bastion of full governmental control over agricultural trade, have been significantly curtailed. Future markets in agricultural products - earlier banned under various statutory orders - are now permitted in several commodities. In 2003, the Model Market Act was passed to reform the regulatory nature of agricultural markets and to allow the private sector to establish parallel markets for the agricultural commodities. The same Act also allows entry of corporate sector in agriculture through ‘Contract Farming’. If these regulatory changes have had the intended impact, integration of

[1] This is part of a wider study of Indian agricultural markets conducted collaboratively between the Asian Economics Centre, University of Melbourne and the National Council for Applied Economics Research, New Delhi, with financial support from the Australian Centre for international Agricultural Research (ACIAR) also involving researchers from several other institutions.
[2] See Jha et al (2005) for a review .

internal (regional) markets and the integration of domestic markets with th world market should have improved since 1994.

In the context of India, the issue of market integration is a central policy issue with major economic and political implications. First, market integration is closely linked to food security. Indian food grain production is spatially diversified, and national food consumption requires substantial inter-regional trade between surplus and deficit regions. The capacity to ensure that food requirements of deficit regions can be met in timely fashion is an essential requirement for Indian food security. If food supplies can be brought in quickly in response to price signals, high price spikes are eliminated and consumption, particularly of the poorer consumers, can be prevented from undesirable falls. In this sense, both improved internal market integration and access to global markets improves food security. Secondly, without improved market integration, other potential welfare gains from market liberalisation cannot be fully captured by agricultural producers and consumers. Thirdly, India's fulfilment of market access requirements under international agreements depends on the extent to which trade liberalisation at the border is translated into price changes within the country.

As a result of foreign trade liberalization, exports of rice started increasing from the mid nineties (Table 1). Following the measures announced in 1994 to liberalise international trade in rice trade, exports of almost all major agricultural commodities have been liberalized. Licensing arrangements have been relaxed, tariffs have been reduced, many items have been freed from quantitative restrictions, and the private sector has been permitted to import most food items. The general trend has been towards lower tariffs, though domestic political pressures have at times reversed this process. The tariff rates were reduced sharply over the decade from a weighted average of 72.5% in 1991-92 to 24.6 in 1996-97, but rose again in the late nineties to 35.1% in 2001-02.[3] In this context, it seems reasonable to expect that linkages between the Indian domestic market and world markets would have strengthened.

3. Spatial Market Integration

[3] In 2000-01 tariffs were raised allegedly to counter possible dumping (Bathla, 2006).

The extent to which spatially separated markets become integrated depend on trade costs (reflecting both trade barriers across relevant spatial boundaries and transport costs) as well as on market structures. If spatially separated markets are linked by trade, and prices always differ only by unit transport costs, then markets are spatially integrated. If transport costs do not change, price movements in perfectly integrated markets will be identical. If domestic and international markets are integrated, international price changes will be fully transmitted to domestic markets provided transport costs and other trade costs remain unchanged. In practice market integration is a dynamic long run process, with prices in integrated markets tending to converge through trade related short run adjustments, with the speed of convergence depending on market structures and frictions.[4] Following the market liberalization and structural adjustment policies undertaken by a number of developing countries in the recent period, the degree to which markets are integrated has been used quite extensively as a yardstick in assessing the success of policy reforms, (see, e.g., Alexander and Wyeth (1994), Baulch (1997), Dercon (1995), Goletti and Babu (1994), Gordon (1994)).[5]

In much of the literature, the focus tends to be on trade costs and, in the context of policy liberalization, on changes in the trade regime affecting trade costs at the 'borders'. However, it should be emphasized that market structure can be a major actor: non-competitive market structures can severely inhibit spatial market integration. The link between trade liberalization at the border and internal market structures can be illustrated by considering a simple case of a domestic firm has monopoly rights over imports and internal distribution. For simplicity, we will also assume that there is no domestic supply, though this assumption can be easily relaxed, and that the country is a price taker in world markets. This means that the relevant marginal cost of imports is the exogenously given world price. Let us start with the case where there is an *ad valorem* tariff 't' on imports, and the world price of $P_{w.}$ If the import monopoly is a profit maximiser, it will equate marginal cost to marginal revenue and set the domestic sale price P_t higher than the $(1+t)\ P_w$, the price at which it can import, at as shown in figure 1. Suppose there is trade liberalization and the tariff is removed. The price facing the firm falls by the full amount of the tariff, but

[4] For a review of key concepts in spatial market integration, see Ravallion (1986).
[5] However, note that market integration by itself does not imply an efficient spatial allocation.

the firm sets its domestic sale price at P_f. If the domestic market was fully competitive, the domestic price would have been (1+t) P_w before trade liberalization, imports Q_t and price P_w and imports Q_f afterwards. As can be seen, the fall in domestic price and increase in imports is lower than would have occurred if the domestic market structure was competitive. In reality, preferential treatment in foreign and domestic trade is often granted – as was the case in India - to state trading enterprises, who may not be simple profit maximizing firms, and changes following liberalization may be somewhat different. [6]

There are several recent studies of overall spatial market integration within India based on analysis of consumer price indices in various locations (Das and Bhattacharya (2004), Virmani and Mittal (2006) as well as several studies of internal agricultural market integration in India which have indicated considerable imperfections due to several distortions and government interventions (for recent studies, see Kumar (2006), Jha et. al. (1997, 2005), Kumar and Sharma (2003), Wilson (2001)). Even the most recent of these studies (Jha et al, 2005) concludes that Indian agricultural markets remain highly segmented; the implication is clealy that recent reforms have had no major impact.

However, these studies have two major methodological limitations that constrain their capacity to shed light on recent developments in the Indian rice market and the impact of reforms. First, they have focused purely on domestic spatial market integration, and have not ignored the fact that since 1994 international trade in rice was liberalised allowing the Indian rice market to move towards integration with world markets Secondly, existing studies all implicitly assume that the fundamental structure of market integration has remained unaltered over time and examine market integration using data drawn from periods that encompass both pre- and post- policy reform periods. Of course , in the context of the reform process, what is really important is to discover *if market integration is changing - improving - over time, both within India*

[6] These are discussed in several papers by McCorriston, and MacLaren (for example, 2005a, 2005b)

and with international markets, rather than that integration was absent or weak in the past which is primarily only of academic interest.[7]

In this paper, we aim to overcome those two limitations and focus on whether policy reforms have improved market integration since the major reforms in 1994.

4. Methodology

There have been many econometric methods used in the studies of market integration, with the early ones being based on bivariate correlations of price series in different markets. However, since Ravallion (1986), dynamic models that overcame several limitations of the simple bivariate approaches have become popular. A further methodological advance came with the development of multiple cointegration method of Johansen (1988), and its application has since become standard in the studies of market integration. However, an important limitation with this approach is that it is not capable of fully utilising the information available when a relatively large number of regional units or locations exist, as it only permits analyses of a relatively small number of markets.

In the context of multiple cointegration, Gonzales-Rivera and Helfand (2001) considered a market with n geographically distinct locations. They defined the market integration as the case where the prices of n locations are cointegrated with (n-1) cointegrating vectors or, equivalently, with one common factor. They have adopted a sequential testing procedure to identify a set of locations that are cointegrated with one common factor, using the trace statistic of Johansen (1988). They found that 15 locations in the Brazilian rice market are cointegrated with one common factor. Based on the same method, Jha et al. (2005) examined the case of Indian rice market. However, as Gonzales-Rivera and Helfand (2001, p579) noted, the statistical validity of their sequential testing procedure is questionable. In addition, the dimensionality problem of the VAR model can substantially undermine the performance of their test, especially when the test is performed with a number of locations as large as 15.

[7] Virmani and Mittal (2006) have compared estimates of spatial variability in prices in domestic markets for several commodity categories at two discrete points in time (in 1994 and 2004) and concluded that market integration seems to be better in 2004.

Given this potential shortcomings of the multiple cointegraton method in a large VAR system, this paper takes a different approach. We test whether the prices in different locations are convergent over time, by utilising the panel unit root tests which are designed to exploit larger and richer data sets (see for example, Abauf and Jorion, 1990). Although the panel unit root test is useful as a mean of testing for the convergence of a set of time series, Maddala and Kim (1998; p.138) argued that it is of limited value in practice because it does not reveal the speed of convergence of individual time series. In this paper, we also estimate the half-life of convergence to measure the speed of adjustment in price differentials, which allow us to address the important policy issue of how policy reforms of recent years have affected market integration in India.

4.1 Methodology: Panel Unit Root Tests

As mentioned above, an attraction of panel unit root testing is that, by pooling the observations from different cross-sectional units, the test can enjoy a larger sample size, which can give rise to a higher power (see, for a recent review, Breitung and Pesaran, 2005). In this section, we provide brief descriptions of the panel unit tests used in this study.

4. Methodology

In this section, we provide brief descriptions of the panel unit tests and the method of half-life estimation adopted in this study.

4.1 The Im-Pesaran-Shin (IPS) test

Im et al. (1997) considered the model of the form

$$Y_{it} = \alpha_i + \beta_i t + z_{it}; \qquad \Delta z_{it} = \phi_i z_{it-1} + \sum_{j=1}^{p} \gamma_{ij} \Delta z_{it-j} + e_{it} \ ,$$

where i (= 1, …., N) indicates a cross-sectional unit, t (= 1, …, T) is a time index and e_{it} ~ IID(0,σ^2). They specified the null and alternative hypotheses of the form

$$\begin{aligned} &H_0 : \phi_1 = \phi_2 = ... = \phi_N = 0 \\ &H_1 : \phi_1 < 0, \phi_2 < 0, ..., \phi_{N_0} < 0 \qquad (N_0 \le N) \end{aligned} \tag{1}$$

The null hypothesis indicates that all time series in each cross-sectional unit are non-stationary with a unit root. Under the alternative hypothesis, at least N_0 time series are stationary. The test statistic is constructed from the t-test statistics calculated from individual cross-sectional units. Let $\hat{\tau}_i$ denote the augmented Dicky-Fuller (ADF) t-statistic to test for $\phi_i = 0$. Im et al. (1997) have shown that

$$\frac{\sqrt{N}(\hat{\tau}_N - E(\hat{\tau}_i))}{\sqrt{Var(\hat{\tau}_i)}} \Rightarrow N(0,1),$$

where $\hat{\tau}_N = \frac{1}{N}\sum_{i=1}^{N}\hat{\tau}_i$. That is, the average of $\hat{\tau}_i$ statistics over all cross-sectional units converges to the standard normal distribution, when appropriately standardized. The values of $E(\hat{\tau}_i)$ and $Var(\hat{\tau}_i)$ are tabulated in Im et al. (1997).

4.2 Fisher test

Maddala and Wu (1999) and Choi (2001) suggested a test for the null and alternative hypotheses given in (1) based on the p-values of individual statistics, which is an approach originally proposed by Fisher (1932). Let π_i denote the p-value of the individual t-statistic $\hat{\tau}_i$. According to Fisher (1932), the statistic $-2\sum_{i=1}^{N}\log(\pi_i)$ follows the chi-squared distribution with $2N$ degrees of freedom. Alternatively, Choi (2001) has shown that

$$Z = \frac{1}{N}\sum_{i=1}^{N}\Phi^{-1}(\pi_i) \Rightarrow N(0,1),$$

where Φ is the standard normal cumulative distribution function.

4.3 Half-life estimation

The half-life, defined as the number of periods required for the impulse response to a unit shock to a time series to dissipate by half, is widely used as a measure of persistence in economic time series. [8]

[8] It is particularly important in the context of testing for the validity of parity conditions in international economics. For example, mean-reversion of real exchange rates is a key condition for the empirical validity of purchasing power parity (Rogoff, 1996).

Model and estimation

The half-life is often estimated from the autoregressive (AR) model of the form

$$Y_t = \mu + \beta t + \alpha_1 Y_{t-1} + \ldots + \alpha_p Y_{t-p} + u_t, \qquad (2)$$

where u_t ~ iid $(0,\sigma^2)$. Note that we suppress the subscript i for a cross-sectional unit for notational simplicity. The AR model given in (2) can be expressed as an MA(∞) model with the coefficients $\{\psi_i\}_{i=0}^{\infty}$ where ψ_0 = 1 and ψ_i represents the impulse response of Y_{t+i} to a unit shock in u_t at time t, i.e. $\psi_i = \partial Y_{t+i}/\partial u_t$, for i = 0, 1, 2, … . The plot of $\{\psi_i\}_{i=0}^{m}$ against i, for a reasonably large integer m, is called the impulse response function of Y, which describes how a time series responds to a unit shock in the error term over a time period of length m. The half-life h is calculated as the largest value j which satisfies $|\psi_{j-1}| \geq 0.5$ and $|\psi_j| < 0.5$. A closed form solution exists in the AR(1) case, i.e., $h = \log(0.5)/\log(\alpha)$. For an AR($p$) model with $p > 1$, the value of h can be obtained from $\{\psi_i\}_{i=0}^{m}$. When j is a number between i-1 and i, linear interpolation is used to determine the value of h.

Given the observed time series $\{Y_t\}_{t=1}^{n}$, the least-squares (LS) estimator for $\gamma = (\mu, \beta, \alpha_1,\ldots,\alpha_p)$ in equation (1) can be obtained by regressing Y_t on (1, t, $Y_{t-1}, \ldots, Y_{t-p}$). The LS estimator and the associated residuals are denoted as $\hat{\gamma} = (\hat{\mu}, \hat{\beta}, \hat{\alpha}_1, ..., \hat{\alpha}_p)$ and $\{\hat{u}_t\}_{t=p+1}^{n}$ respectively. In the AR(1) case, the half-life is estimated as

$$\hat{h} = \begin{cases} \log(0.5)/\log(\hat{\alpha}_1) & \text{if } \hat{\alpha}_1 < 1 \\ \infty & \text{otherwise} \end{cases}.$$

For a higher order model, $\hat{h}$ is obtained from the estimated impulse response function $\{\hat{\psi}_i\}_{i=1}^{m}$, where $\hat{\psi}_i$ is the ith coefficient in the MA(∞) representation associated with $\hat{\gamma}$. When the model has a characteristic root close to one, $\hat{h}$ may not be found even with a reasonably large value of m, since $\{\hat{\psi}_i\}_{i=1}^{m}$ declines fairly slowly. In this case, we use an approximation

$$\hat{h} = \begin{cases} \log(0.5)/\log(\hat{\alpha}) & \text{if } \hat{\alpha} < 1 \\ \infty & \text{otherwise} \end{cases},$$

where $\hat{\alpha} = \hat{\alpha}_1 + ... + \hat{\alpha}_p$, following Murray and Papell (2002). In this paper, we set m = n and use this approximation if $\{\hat{\psi}_i\}_{i=1}^{n}$ does not reach 0.5 for $i \le n$.

Bias-corrected bootstrap for point and interval estimation

The above procedures describe point estimation of the half-life. However, provision of only a point estimate is not informative, since the half-life can take any value between 0 and infinity. In addition, since it takes a ratio form, the half-life estimator $\hat{h}$ is biased in small samples, and it has unknown (sampling) distributional properties with possibly non-existent finite sample moments. Given these properties, a number of past studies proposed the use of the bootstrap method for interval estimation, with a built-in bias-correction procedure; see Murray and Papell (2002, 2005), Caporale et al. (2005), Rapach and Wohar (2004), Gospodinov (2004) and Rossi (2005). However, the methods advocated by these authors frequently provide confidence intervals whose upper bounds are infinite, even though the underlying time series is stationary and convergent, as demonstrated in a Monte Carlo study conducted by Kim et al (2007).

Kim et al. (2007) proposed an alternative bias-corrected bootstrap procedure, in which the highest density region (HDR) method of Hyndman (1996) is used to construct bias-corrected point estimator and confidence interval. The HDR method provides a more sensible way of point and interval estimation than the conventional methods for half-life estimation. Their Monte Carlo experiment revealed that their bias-corrected bootstrap HDR confidence interval provides much tighter and more informative confidence interval for half-life with enhanced coverage properties. In addition, it is found that the HDR point estimator also performs better than other bias-corrected point estimators.

We provide a brief description of the bias-corrected bootstrap procedure proposed by Kim et al. (2007). First, we obtain the bias-corrected version of $\hat{\gamma} = (\hat{\mu}, \hat{\beta}, \hat{\alpha}_1, ..., \hat{\alpha}_p)$ using the non-parametric bootstrap. Generate a pseudo-data set $\{Y_t^*\}_{t=1}^{n}$ as

$$Y_t^* = \hat{\mu} + \hat{\beta}t + \hat{\alpha}_1 Y_{t-1}^* + ... + \hat{\alpha}_p Y_{t-p}^* + e_t^*, \quad (3)$$

using $\{Y_t\}_{t=1}^{p}$ as starting values, where e_t^* is a random draw with replacement from $\{\hat{u}_t\}_{t=p+1}^{n}$. The above process can be repeated many times so that B_1 sets of pseudo-data are generated, from which B_1 sets of bootstrap parameter estimates for γ, denoted $\{\gamma^*(j)\}_{j=1}^{B_1}$, can be obtained. A typical γ^* = (μ^*,β^*, $\alpha_1{}^*$,…,$\alpha_p{}^*$) is obtained by regressing Y_t^* on (1, t, Y_{t-1}^*,...,Y_{t-p}^*). The bias of $\hat{\gamma}$ can be estimated as $Bias(\hat{\gamma})$ = $\bar{\gamma}^* - \hat{\gamma}$, where $\bar{\gamma}^*$ is the sample mean of $\{\gamma^*(j)\}_{j=1}^{B}$. The bias-corrected estimator $\hat{\gamma}_B^c = (\hat{\mu}^c, \hat{\beta}^c, \hat{\alpha}_1^c,..., \hat{\alpha}_p^c)$ for γ can be calculated as $\hat{\gamma} - Bias(\hat{\gamma})$.

To obtain the bias-corrected point and interval estimators for half-life, we conduct the second-stage bootstrap using the bias-corrected parameter estimators obtained above, following the bootstrap-after-bootstrap of Kilian (1998). Generate the pseudo-data set $\{Y_t^*\}_{t=1}^{n}$ recursively as

$$Y_t^* = \mu^c + \beta^c + \alpha_1^c Y_{t-1}^* + ... + \alpha_p^c Y_{t-p}^* + v_t^*, \tag{4}$$

using $\{Y_t\}_{t=1}^{p}$ as starting values, where v_t^* is a random draw with replacement from $\{u_t^c\}_{t=p+1}^{n}$. Using $\{Y_t^*\}_{t=1}^{n}$, the parameters of the AR(p) model are estimated with bias-correction to obtain $(\mu^{c*}, \beta^{c*}, \alpha_1^{c*},..., \alpha_p^{c*})$. The associated half-life estimate is denoted as h^*. Repeat (4) and estimation of h^* many times, say B_2, to obtain the bootstrap-based distribution of the half-life estimates $\{h_i^*\}_{i=1}^{B_2}$.

To obtain a tight and informative confidence interval from $\{h_i^*\}_{i=1}^{B}$, Kim et al. (2007) used the HDR method of Hyndman (1996). Let $f(x)$ be the density function for a random variable X. The 100(1-θ)% HDR is defined (Hyndman, 1996) as the subset $R(f_\theta)$ of the sample space of X such that $R(f_\theta) = \{x: f(x) \geq f_\theta\}$, where f_θ is the largest constant such that $\Pr[X \in R(f_\theta)] \geq 1 - \theta$. $R(f_\theta)$ represents the smallest region with a given probability content. In short, the HDR method produces confidence intervals concentrated around the mode of the distribution. In the present context, X is the half-life estimator of a time series and its density can be estimated from the bootstrap replicates of the half-life $\{h_i^*\}_{i=1}^{B_2}$. We estimate the density $f(x)$ using a kernel estimator with the Gaussian kernel, with bandwidth selected using the Sheather-Jones (1991) rule. From the estimated density, the mode of the distribution is used as the bias-corrected point estimator for the half-life, along with the interval concentrated around the mode of the distribution with the probability content $100(1-\theta)\%$. In the multi-modal case, the global mode and the associated interval are used as point and interval estimates of half-life.

5. Data

We use monthly data for rice prices from 1980:4 to 2002:12 (273 observations) for 23 cities in India. Most importantly, in light of policy reforms that have been designed to integrate the previously almost completely insulated Indian domestic market with the international rice market, we also include the international rice price as an extra cross-sectional unit, resulting in 24 cross-sectional units in total. Monthly wholesale prices of the above mentioned states/markets are taken from the Ministry of Agriculture, Government of India, documents – ‘Agricultural Prices in India’, and ‘Agricultural Situation in India’. The monthly international prices are taken from the IMF yearbook for the Thailand (5%) broken rice price. All the prices (domestic as well as international) are converted into US$ and have been used as natural logarithms in the model. The price ratio in natural logarithm is used to measure the convergence. That is, $Y_{it} = \log(P_{it}/P_{0t})$, where P_{it} is the rice price for ith cross-sectional unit at time t and P_{0t} is the rice price for the numeraire at time t.

In keeping with the focus of our investigation, we divide the observations into two period, Period 1 (pre-reform period, observations up to 94:12), and Period 2 (post-reform period), and use the international price as the numeraire. [9]

6. Results

Table 2 presents the panel unit root results. For both periods, the null hypothesis of unit root is soundly rejected at 1% level of significance, according to both the IPS and Choi’s Z tests, indicating that the price differentials are overall convergent. Notice

[9] In ongoing work, we are using the unit root test proposed by Westerlund (2006) which allows for multiple unknown breaks to investigate structural breaks in the data in greater depth.

that the p-values (both from the panel unit root and individual unit root tests) in Period I are larger than those in Period II. Although its statistical justification may be arguable, this observation suggests that the prices differentials in Period II converge at a faster rate than those of Period I. In order to substantiate this claim in a more precise manner, we turn to our half-life estimates for the two periods, using the international price as the numeraire.

The results are presented in Table 3. Note that the half-life estimates are expressed in years. The table reports bias-corrected bootstrap HDR point estimates, as well as their 90% confidence intervals, along with the estimates of the persistence parameter (alpha: the sum of all AR slope coefficients). The results clearly indicate that the speed of convergence in the post-liberalisation period improved remarkably. On average, the value of half-life point estimates declined from 4 years to less than half year. Not even a single market observed any increase in convergence time from the pre to post liberalization period. In no case is the point estimate higher than a half-year in the post-reform period. Moreover, the confidence intervals are much tighter in Period II, further strengthening the case for faster rate of convergence. [10]

What influences the variations in convergence speeds in different markets? In a bid to explain the convergence speeds, we tried to relate the half-life estimates in different markets with city/state level variables. We hypothesised that distance of markets centres from the nearest port where from foreign trade can take place would be a critical variable influencing price transmission between foreign and domestic markets. In addition, the level of infrastructure and nature of market structure are likely to influence the speed of price adjustments through trade, both with world markets and other domestic markets. An infrastructure index that measures infrastructure quality is available for various Indian states and we used variable reflecting the existence of infrastructure and we used this infrastructure index for the period 2000-01 as an

[10] The very high point estimates for three cities – Nizammabad, Bangalore and Sambalpur (with lower bounds of the 10% confidence intervals over 13 years) suggest that price series of these cities were practically non-convergent with international prices in the pre-1994 period. Even if these estimates are treated as outliers, the mean point estimate of the remaining cities for pre-liberalisation period is more than double the mean for the post- liberalisation period.

explanatory variable.[11] Before the liberalisation reforms, state interventions were pervasive and internal markets were heavily distorted. However, states that have implemented market reforms have moved towards more competitive market structures. But there is no single measure that can capture the nature of market structure in each location. We experimented with several variables to proxy market structure, such as whether some major domestic market reforms have been undertaken by different states [12], market arrivals as a % of production which reflects the market orientation of rice agriculture in the state, and a 'market intervention' variable that measures the degree of market interventions by the Food Corporation of India (which intervenes in the rice markets purchasing from the farmers at the state declared Minimum Support Price (MSP) to maintain a floor price). The market surplus variable is a measure of the market orientation of the rice industry and, given the importance of rice as a consumption staple in all households, also indicative of the net supply situation of the state. The results of the regression models with distance, infrastructure and different market structure proxies for the post-reform period are shown in table 4.

The variable on distance from the port did not turn out significant in any of the equations. However, the infrastructure variable was generally significant (though at a lower probability level). The infrastructure index is a composite that captures the overall quality of the state's infrastructure facilities such as road and rail connectivity, better communication facilities etc. It appears that the overall quality of infrastructure is a more important factor than the physical distance to the nearest port. The market reform variable was not significant possibly because the process of market reforms has started only very recently, after the Central Government passed the Model Market Act in 2003. Somewhat surprisingly, the market intervention variable – reflecting interventions by the FCI turned out to be significant and positively influencing the speed of convergence (lowering the time taken for convergence).[13] One interpretation

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[11] Ideally we would have liked to use values of this index for each year but the complete time series for the index was not available to us for the entire period. We hope to extend the analysis with annual values of the index in future work.

[12] The main criterion for assessing the progress of pro-competitive legislative changes was whether the states had passed legislation fundamentally changing the provisions of the Agriculture Produce Marketing Committee (APMC) Act. The APMC Act was the most important regulatory measure that governed agricultural markets under which no private trader or farmer was allowed to engage in trade except through state regulated markets.

[13] A similar result was obtained by Bathla (2006).

of this result is that market interventions to support the minimum support price are facilitating market integration. However, it is more likely that this variable was capturing the impact of large market surpluses, as it is in high surplus states that FCI interventions have been most prominent. When the regression models including the market surplus variable were estimated, this variable became highly significant with a negative sign, indicating that time for convergence was lower in those states that had higher market surplus. This is consistent with the position of India as a net rice exporting country in recent years, and suggests that liberalising rice exports has indeed facilitated faster convergence of internal prices with international prices in rice surplus states.

7. Conclusions

We have analysed the behaviour of spatial rice prices in a number of cities scattered across India to investigate whether agricultural policy reforms in India have improved spatial market integration both internally and with world markets. Our results suggest that policy liberalisation, particularly the major reforms of 1994 that liberalised rice exports, appear to have significantly improved market integration. This conclusion contrasts sharply with recent analyses of such market integration that have concluded that Indian rice markets continue to be quite fragmented and segmented. However, foreign trade liberalisation is not symmetric in the way it deals with exports and imports, and this has implications for the spatial patterns of price convergence. Reflecting the dominance of producer interests in policy formation, exporting has been made significantly more liberal than importing. As a result, surplus locations converge faster to international prices. Better infrastructure, as expected, facilitates faster price convergence. It is too early to measure the impact of recent changes to domestic market structures brought about as a result of amendments to the APMC act, but these are likely to improve market integration further. Overall, we conclude that policy liberalisation has significantly improved market integration, and this process would accelerate further with expected improvements to infrastructure and ongoing domestic market reforms.

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Figure 1: Trade Liberalisation and Domestic Price: the case of a profit maximising import monopoly

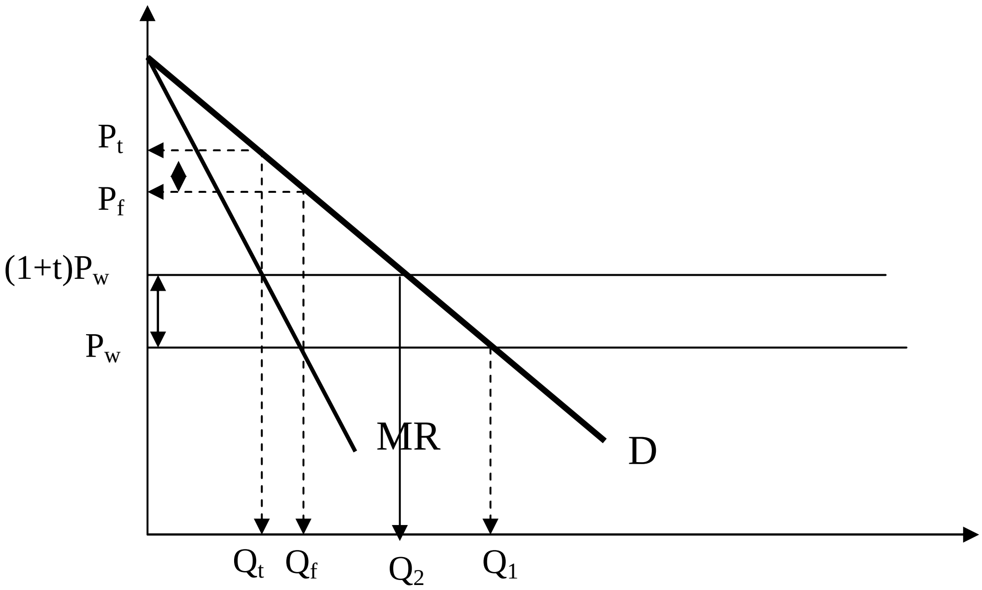

Table 1: Production and distribution of rice in India

	Production (MT)	As a percentage of production				
		Market Arrivals	Procurement	Off-take	Exports	Imports
1980-81	53.6	30.20	11.56	10.14	1.35	0.01
1985-86	63.8	31.20	14.32	12.03	0.38	0.10
1990-91	74.3	30.90	15.72	13.11	0.68	0.09
1995-96	77.0	39.70	13.08	12.29	6.38	0.00
1996-97	81.7	42.30	15.87	13.63	3.07	0.00
1997-98	82.5	41.10	18.89	11.99	2.91	0.00
1998-99	86.1	39.00	14.64	12.48	5.77	0.01
1999-00	89.5	44.50	20.37	12.64	2.04	0.04
2000-01	85.0	Na	25.04	9.38	1.81	0.02
2001-02	93.3	Na	23.72	8.75	2.37	0.00
2002-03	71.8	Na	22.87	10.29	6.92	0.00
2003-04	88.2	Na	25.88	na	3.87	0.00
2004-05	85.3	Na	28.93	na	5.62	0.00

Source: Based on data from various issues of Government of India, ‘Agricultural Statistics at a Glance’, ‘Economic Survey’ and ‘Bulletin on Food Statistics’

Table 2: Unit Root Tests (International price as numeraire)

Panel Unit Root Test	**Period I**	**Period II**
IPS test	0.0031	0.0000
ADF Fisher (Choi-Z) Test	0.0027	0.0000
Individual (ADF) test		
Vijaywada	0.1220	0.0442
Gauhati	0.2625	0.0551
Patna	0.3368	0.3647
Amrtsr_rice	0.0379	0.0908
Simoga	0.2505	0.1978
Nagpur	0.3972	0.0561
Tirunelveli	0.1049	0.1232
Kanpur	0.5389	0.1055
Sainthia	0.3066	0.1320
Kakinada	0.2519	0.0048
Nizamabad	0.4049	0.0053
Ranchi	0.2586	0.4030
Dumka	0.2461	0.3668
Arrah	0.4331	0.0624
Gaya	0.2756	0.0869
Bangalore	0.5117	0.0774
Trivendrum	0.4978	0.6565
Sambalpur	0.4027	0.2120
Cuttack	0.3187	0.3204
Allahabad	0.2585	0.3114
Contai	0.1851	0.0785
Siliguri	0.1831	0.1985
Delhi	0.2812	0.0179

Notes:
The entries are the p-values of the test.
Period I is to 1980:04 to 1994:12, and Period II is 1995:1 to 2002:12

Table 3: Half Life Estimates (International price as the numeraire)

	Period I				Period II			
	alpha	**Point**	**90% CI**		**alpha**	**Point**	**90% CI**	
Vijaywada	0.92	0.79	0.22	5.36	0.81	0.33	0.08	3.00
Gauhati	0.92	0.65	0.23	5.59	0.78	0.27	0.07	2.25
Patna	0.94	0.76	0.22	6.39	0.87	0.40	0.10	3.76
Amrtsr_rice	0.87	0.39	0.14	3.64	0.86	0.34	0.10	3.25
Simoga	0.93	0.86	0.22	7.24	0.88	0.43	0.15	3.84
Nagpur	0.92	0.75	0.22	6.77	0.82	0.34	0.11	3.38
Tirunelveli	0.91	0.59	0.18	6.21	0.84	0.32	0.09	3.34
Kanpur	0.95	1.02	0.28	8.91	0.71	0.21	0.09	0.97
Sainthia	0.93	0.96	0.23	6.31	0.85	0.40	0.13	3.05
Kakinada	0.94	0.87	0.21	6.50	0.77	0.25	0.10	2.39
Nizamabad	0.95	23.92	14.36	146.82	0.75	0.24	0.06	1.92
Ranchi	0.94	0.71	0.24	8.00	0.81	0.31	0.11	2.48
Dumka	0.94	1.02	0.27	7.95	0.89	0.45	0.15	4.25
Arrah	0.93	1.18	0.27	6.07	0.81	0.29	0.08	3.32
Gaya	0.93	0.87	0.25	6.89	0.87	0.45	0.10	3.54
Bangalore	0.96	28.91	15.24	174.10	0.83	0.31	0.09	3.26
Trivendrum	0.96	1.02	0.32	9.27	0.93	0.49	0.11	4.27
Sambalpur	0.94	22.86	13.59	117.35	0.89	0.41	0.12	3.88
Cuttack	0.95	0.91	0.23	7.62	0.90	0.45	0.11	3.60
Allahabad	0.94	0.79	0.22	7.20	0.88	0.48	0.16	4.67
Contai	0.93	0.84	0.26	7.97	0.85	0.35	0.13	3.52
Siliguri	0.94	1.09	0.28	6.48	0.86	0.37	0.10	3.68
Delhi	0.93	0.76	0.26	8.49	0.78	0.23	0.06	2.92
Mean	0.93	4.02	2.08	25.09	0.84	0.35	0.11	3.24

Notes:

1. alpha is the sum of AR coefficients.
2. Point is the HDR point estimate in years. 90% CI for the half life is expressed in years.
3. Period I is to 1980:04 to 1994:12, and Period II is 1995:1 to 2002:12
4. For Period I, AR(1) models are fitted for all cases, except for Vijaywada, Gauhati, Kanpur, Ranchi, Dumka, Arrah, Gaya, Trivendrum, Contai, Siliguri, and Delhi, to which AR(2) models are fitted.
5. For Period II, AR(1) models are fitted for all cases, except for Simoga , Nagpur, Kanpur, Sainthia, Kakinada, Ranchi, Dumka, Allahabad and Contai, to which AR(2) models are fitted.

Table 4: Regression results explaining half-life estimates in the post reform period –state-wise

	Model 1	**Model 2**	**Model 3**	**Model 4**
Constant	0.313 (4.9)	0.298 (4.4)	0.362 (6.1)	0.361 (5.9)
Distance	-0.000 (-1.6)	-0.000 (-1.5)	-0.000 (-0.04)	-0.000 (-0.04)
Infrastructure	0.001 (1.5)	0.001 (1.6)	0.001 (1.3)	0.001 (1.3)
Market intervention (% of FCI purchases of market arrivals)	-0.001 (-1.9)	-0.002 (-1.9)	-	-
Market Reforms (Dummy: reforms=1)	-	0.034 (0.8)	-	0.006 (0.16)
Market Surplus (market arrivals as % of aggregate rice production)	-	-	-0.002 (-2.5)	-0.002 (-2.3)
R^2	0.29	0.31	0.32	.032
F-Statistic	2.59	2.05	2.65	1.88

Note: Figures in parentheses are respective ‘t’ values.