Integration of the International Rice Market: Implications for Trade Liberalization

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April 2002

ABSTRACT

A cointegration analysis performed in this paper finds that the international rice market is highly segmented. This is explained by the high degree of market intervention in this sector due to food security reasons and by strong consumer preferences for specific rice varieties. This finding is supported by subsequent estimation of the elasticities of substitution between rice imported from different sources. The estimates are found to be substantially lower than those often used in e.g. computable general equilibrium models. This implies that trade liberalization will have less of an impact than otherwise predicted.

Key words: cointegration analysis, estimation of Armington elasticities, international rice trade

Introduction

Computable general equilibrium (CGE) models form the basis for much applied research on trade issues, hereunder the impacts of trade liberalization. As with any empirical economic modeling framework, a CGE analysis is no better than the theory and data underpinning it. A critique that has been raised against some CGE models is the lack of econometric foundation for the behavioral parameters used in the models (e.g. McKitrick). Furthermore, many of the elasticity estimates being used are fairly dated. In a global CGE model the parameters in question typically include the source elasticities, the factor substitution elasticities, the factor transformation elasticities, the investment parameters, and the consumer demand elasticities.

This paper contributes to the continued efforts to supply CGE models with a more solid econometric foundation. Our efforts focus on one type of elasticity – the elasticity of substitution among imports from different sources used in the well-known Armington aggregation structure, and one commodity – rice. The literature on elasticities of substitution between import sources is fairly scarce – most of the econometric evidence available is related to domestic-import substitution. Yet terms of trade effects in a global CGE analysis will depend heavily on these elasticities. If they are high, a given region can displace other regions in its export markets without lowering its export prices very much. Conversely, if elasticities are low, a given importer will be sensitive to changes in a particular exporter's capacity to supply the foreign market.

* The authors would like to thank Kim M. Lind (Danish Research Institute of Food Economics), Hans Christian Kongsted and Heino Bohn Nielsen (University of Copenhagen) for useful discussions and suggestions regarding data and methodology, and Mark Gehlhar (ERS/USDA) for providing part of the data used in this study.

Before estimating these elasticities one has to clarify whether or not the Armington structure is appropriate for all segments of the market. Cointegration analysis is therefore used to identify whether certain markets are integrated to an extent that warrants treating them as homogenous goods rather than as heterogenous goods as assumed in the Armington structure. The results of the cointegration analysis then provide a basis upon which the elasticities of substitution between different sources are estimated for several major importers of rice. Finally, these estimates are incorporated into a fairly standard global CGE model (the Global Trade Analysis Project (GTAP) model, see e.g. Hertel) to illustrate the importance of using econometrically founded parameters when performing trade analysis. Before embarking on the formal econometric analyses, a brief introduction to the characteristics of the international rice market is given in the next section.

International rice trade: structure and policy instruments[1]

The international market for rice is often characterized as a thin and volatile residual market with elements of instability and uncertainty that distinguish it from world markets for wheat and maize (Barker and Herdt, Latham). The world rice market is thin in the sense that the amount of rice traded internationally is small relative to total production. Over the period 1961-99 the average trade-in-production shares have been 18.2% for wheat, 13.7% for maize and only 4.6% for rice (FAOSTAT). This small trade share implies that fluctuations in production have magnified effects on traded volumes.

As shall be discussed shortly, however, government rice policies clearly contribute to the thinness and instability of the global rice market. Combined with the inevitable variations in harvest performance, tight controls of exports and imports, for example, mean that both the level and sources of demand and supply of rice are rather unpredictable. Thinness in itself is not necessarily a problem if sellers and buyers are the same each year. Yet the international rice market is extremely volatile in that sellers and buyers enter the market at a given point in time depending on the performance of their own domestic crop (Latham).

Rice exports are concentrated in the hands of just a few large exporting nations. The six largest rice exporters in 1999 were Thailand, Vietnam, China, the USA, India and Pakistan (in volume terms and in that order). Exporting 7 million tons in 1999, Thailand is by far the leading rice exporter, accounting for almost 30% of total world exports (26 million metric tons, according to FAOSTAT). Vietnam ranked second that year, exporting more than 4 million tons, and thereby accounting for 18% of total exports. The United States, China and India each accounted for 10-11% of world exports, and Pakistan settled on 7%. These six exporters have delivered between 73% and 85% of total export volumes over the past 20 years, averaging around 80% in the 1990s.

Rice trade is much more dispersed on the import side than on the export side: in 1999 thirty-five countries made up 80% of total imports. In 1999 the largest single importers were by far Indonesia and Bangladesh. For both countries, import demand in a given year depends on the outcome of the domestic rice crop. Two other significant importers in Asia are the Philippines and Malaysia. The Middle East region is traditionally the world's strongest market for high-quality rice, while the Sub-Saharan African region is a major importer of low quality rice. The Latin American and Caribbean

[1] This section draws heavily on Nielsen's review of national rice policies.

nations accounted for more than 10% of total imports in 1999, with the bulk of it going to Brazil. The importance of the European Union as a rice importer has declined rather substantially over the period 1980-99. From averaging 12.8% of total rice imports in 1980-85, the EU has only purchased 3.9% of world rice supplies during the period 1995-99.

Rice is a staple food in almost all rice producing countries in Asia – the region in which 90% of the world's production and consumption of rice takes place. The primary goals of the rice policies pursued in these countries are therefore self-sufficiency and stable domestic prices. Using a combination of domestic support measures and restrictive trade policy instruments, these countries end up isolating their domestic markets from the world market as they strive to achieve these goals. Hence these isolationist policies simply add to the inherent thinness, unpredictability and instability of the international rice markets. Only concerted dismantling of these trade-distorting policies will change the situation.

In Asia there is a tradition of state trading enterprises (STEs) having strong controls over domestic marketing of rice as well as being given exclusive rights to import and export rice. As the policy review by Nielsen makes clear, there are concerns that importing STEs effectively foreclose domestic markets from foreign competition, whilst exporting STEs distort export competition on world markets. Although only a few countries have explicit rice export subsidies in place, there are other more covert forms of support to exporters, particularly in the US. Such support measures deserve a closer examination in order to assess the extent to which some of them may be classified as export subsidies and hence disciplined under the World Trade Organization (WTO). Other more explicit distortions that affect trade are domestic support measures. Only the developed countries can afford to provide direct support to rice farmers at a significant level. Nevertheless, it is important that this support is reduced because it provides exporters with unfair price competitiveness and raises market prices in importing countries thereby placing foreign suppliers at a disadvantage.

In terms of explicit trade policy instruments, rice imports are subject to a multitude of specific tariffs, combination tariffs, and variable levies (official and de facto). These types of duty provide greater protection than simple *ad valorem* tariffs. Most importantly, they provide increasing protection when world market prices fall. Furthermore, because of this characteristic, the actual degree of protection provided by specific and combination tariffs at a given point in time is not very transparent. Even where *ad valorem* tariffs are used, they are sometimes adjusted so frequently in response to changing domestic and world market conditions that they work like de facto variable levies.

International rice trade is also affected by non-tariff measures such as quantitative restrictions on imports and exports, seasonal bans, tariff rate quotas, etc. As part of the Uruguay Round Agreement on Agriculture several countries introduced Tariff Rate Quotas (TRQs). This mechanism was meant to increase market access, but there are substantial problems related to the use of TRQs. First of all, over-quota tariffs are often prohibitively high. Secondly, the administration of TRQs seems to mean that only selected suppliers benefit from this preferential access. Finally, there were exceptions to the general agreement to introduce TRQs: Japan, Korea and the Philippines, for example, were permitted to use a classic import quota to determine their minimum market access level for rice with no commitments whatsoever to import above this level. The minimum market access commitments made by these three countries have generally been fulfilled, but it may be questioned which effect this has had on international rice trade since the number of countries winning the tenders for these imports has been limited.

Some rice exporting countries benefit from preferential trade agreements. In 1996 the European Union, for example, struck deals with the United States and Thailand on annual tariff rate quotas. The EU notifications to the WTO show that these quotas are virtually filled each year. Moreover, the EU provides preferential conditions and preferential access quotas for rice imports from several developing countries, particularly the African, Caribbean and Pacific (ACP) countries but also India, Pakistan, Egypt and Bangladesh. Around 40% of rice imports into the EU enter on preferential terms (FAS).

Simple conclusions cannot be drawn about the extent to which certain countries' exports are currently being hindered by the rice policies of their (potential) trading partners. In terms of the Japanese and Korean rice markets, for example, there is no doubt that there are severe import restrictions in place – both explicitly and implicitly – and rice exporters have an interest in these being removed. Yet there is also a consumer preference dimension to international rice trade. Demand in these particular markets is primarily for japonica rice varieties whereas lower-income rice exporters such as Vietnam and India are predominantly indica producers. Hence it is not entirely clear to what extent a lifting of these import barriers would in fact boost these countries' export potential. To the extent that farmers are both willing and able to change crop varieties there may be scope for gaining access to such markets in the future as formal trade barriers are dismantled.

Cointegration analysis

Methodology

As mentioned in the introduction a cointegration analysis of the international rice market will determine the extent to which the Armington assumption of product heterogeneity is valid before estimating the elasticities of substitution between import sources. Two spatially distinct markets for the same commodity are said to be integrated if their prices share the same long-term stochastic trend, i.e. the prices cointegrate. If this is so, then price changes in one market will lead to price changes in the other. Factors that affect the degree of market integration include access to price information and distance between markets. Availability of timely and good quality price information promotes market integration. Distance should not in principle be an obstacle to market integration, but it does affect the speed at which arbitrage takes place. Furthermore, it is expected that certain policy interventions such as state trading and government-to-government contracts will affect the degree of market integration as will strong consumer preferences for specific product varieties.

A common technique used to investigate the extent of market integration is cointegration analysis[2] because it allows one to identify whether or not there are stable long-run relations between non-stationary variables. Given that price time series are typically non-stationary and that we are searching for long run stable relations between prices to delineate the international rice market, the cointegration method is therefore highly appropriate. More specifically, cointegration analysis is used in this paper to test whether and for which markets the Law of One Price (LOP) is valid. To the extent that the LOP is

[2] Examples in the field of agricultural and fisheries economics include Zhou, Wan and Chen; Le Goulven; Zanias; Bierlen, Wailes and Cramer; Jaffry et al.; Asche and Hannesson; Goeltti, Ahmed and Farid; and Silvapulle and Jayasuriya.

valid, the Armington assumption of product heterogeneity – which is typically used in global CGE models – is not appropriate. In such a case it would be more relevant to assume product homogeneity and to estimate behavioral relations for the market as a whole rather than distinguished by source. If the LOP is found to be invalid, one can test whether one or several of the price series can be excluded from the analysis. If long run exclusion of a price series is accepted, this implies that the analyzed products are heterogenous and that their markets should be treated separately. If both the LOP test and the long run exclusion test are rejected, the markets are partially coherent, and the products are imperfect substitutes. The underlying idea of this part of the paper is thus to use the two tests (LOP and long run exclusion) to determine the extent to which the international sub-markets for rice should be treated as perfectly coherent (i.e. homogenous products that are perfectly substitutable), completely separate (i.e. non-substitutable goods), or partially coherent (i.e. imperfectly substitutable goods) before estimating the elasticities of substitution between import sources. Given the description of the international rice market in the previous section – particularly the aspects of widespread policy intervention and strong consumer preferences for particular varieties – there is an *a priori* expectation that the international rice market is fairly segmented.

The cointegration analysis framework developed by Johansen takes as its starting point a Vector Autoregressive (VAR) model written in Error Correction Model (ECM) form, i.e.:

(1) $$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k} + \phi D_t + \varepsilon_t, \; t = 1,2,...,T$$

where X_t contains p endogenous variables (in this case prices). The matrix of parameters Π represents the long run solution to the model while the matrices Γ_i contain information about the short run adjustment process. The rank of Π determines the number of stationary linear combinations of the variables in X_t. If Π has reduced rank, i.e. $0 < r < p$, where p is the number of variables, then the matrix can be decomposed into $\Pi = \alpha\beta'$ (where α and β are full rank matrices) and β contains the cointegrating vectors. Cointegration thus implies that while the multivariate series X_t are non-stationary[3], the combination $\beta' X_t$ is stationary.

In determining the rank of Π the Trace test is used. Different distributions for this test apply depending on the composition of the deterministic term ϕD_t and the restrictions imposed hereon. *A priori* one does not expect there to be a deterministic trend in the relationship between the prices since this would imply that the prices would drift from one another in a deterministic manner. Moreover, since it can be argued that there seems to be negative linear trends in the series contained in X_t, the so-called H_1-model (using Johansen's terminology) is used with $\phi D_t = \mu_0$. This model allows for a linear trend in X_t

[3] A prerequisite for the use of cointegration methods is that the data series are integrated of the same order (Note that in this context the word "integrated" or "integration" is an econometric term relating to the properties of time series and has a different meaning from that used in the discussion of integration of markets). The analysis therefore starts by investigating whether the individual data series contain unit roots, i.e. whether they are integrated of order 1, I(1). The regression equation used to perform this test is given as: $\Delta X_t = \pi X_{t-1} + c + \kappa t + \gamma_1 \Delta X_{t-1} + ... + \gamma_{k-1} \Delta X_{t-k+1} + \varepsilon_t$, t = 1,2,...,T. The regression equation contains a deterministic trend because the price series used (described below) are (logs of) real rice prices, which we *a priori* expect to decline over an extended period of time. The graphs of the series support this model. The tests used to determine whether or not the series exhibit non-stationary properties are the Dickey-Fuller F-test and the likelihood ratio test (see e.g. Johansen). The hypothesis is accepted comfortably at the 5% level for all the series using both tests and so it may be concluded that all the series are integrated of order 1. Moreover, there are no signs that the data are I(2) and the series can thus be used in the subsequent cointegration analysis. The test results are presented in Nielsen and Yu.

but no linear trend in $\beta' X_t$. In using the Trace test the null hypothesis is that there are up to r cointegrating vectors. The alternative is that there is exactly one more cointegrating vector. The appropriate distributions for the test are provided in Johansen.

After having determined the number of cointegrating vectors (the rank of Π) the Law of One Price is then tested using Likelihood Ratio tests on β. In the case of two variables that cointegrate the rank will be found to be one and a test of the LOP amounts to testing whether $\beta' = [1,-1]'$. In the multivariate case, the rank will be found to be equal to p-1 and the test of the LOP amounts to testing whether the price series are pair wise cointegrated, i.e. they follow a common trend. For a model with five variables the rank would be found to be 4 and the test on the parameters is

$$(2) \qquad \beta' = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 \\ 1 & 0 & -1 & 0 & 0 \\ 1 & 0 & 0 & -1 & 0 \\ 1 & 0 & 0 & 0 & -1 \end{bmatrix}$$

The test of whether or not the LOP holds amounts to testing whether the restriction imposed on β' makes the linear combination $\beta' X_t$ stationary.

If the rank is less than p-1 the LOP cannot be tested. What can be tested, however, is whether one or more of the series can be excluded from the analysis. Based on the identified rank, long run exclusion can also be tested as a restriction on the matrix β. More specifically, this amounts to testing for a row of zeros in β, or equivalently a column of zeros in β'. In addition to testing whether a particular price series can be excluded from a model, one can also test whether a particular price series can be said to be weakly exogenous. If a price series is weakly exogenous it does not respond to deviations from the long run equilibrium experienced in the previous period. This amounts to testing for rows of zeros in the α matrix in the same way that the test for long run exclusion is performed using the β matrix. This hypothesis is of interest because it allows us to classify markets according to the results of these two tests. Markets may be identified as long run leader markets, long run follower markets, long run segmented markets and long run regulator markets depending on the outcome of these tests. The classification of markets is shown in Table 1.

Data

F.o.b. price series for nine (9) rice types/qualities were obtained from CIRAD[4]: US 2/4, Thai 100, Thai 5% brokens, Vietnamese 5% brokens, Indian 5% brokens, Thai 25% brokens, Vietnamese 25% brokens, Indian 25% brokens, and Thai A1 Super. The price series obtained are on a monthly basis[5]

[4] CIRAD (Centre de coopération internationale en recherche agronomique pour le développement) is a French scientific organization specializing in agricultural research for the tropics and subtropics of the world. The data were obtained at the courtesy of Patricio Mendez del Villar. The data can be made available upon request.

[5] As with virtually all agricultural price series, the rice price series exhibit seasonal patterns. While this seasonality may provide interesting information about the functioning of a market, it may also affect the performance of the cointegration analysis. There are three types of seasonality: deterministic, stationary stochastic and non-stationary stochastic. It is non-stationary stochastic seasonality that may cause problems for the cointegration analysis (see e.g. Engle, Granger and Hallman; and Hylleberg et al.). For the present analysis, the interest lies not so much in the issue of seasonality *per se*, but

and – with the exception of the Indian series – the data series consist of 144 observations from 1990:01 to 2001:12. For the Indian series only 72 observations are available for the period 1996:01 to 2001:12. Each of the series is deflated by the Consumer Price Index (CPI) as a proxy for the general price development in each country.[6] The data are summarized in Table 2.

In the high-quality range, Table 2 reveals that the average price for Thai 100 percent Grade B rice was $213/ton, whilst the price of US long grain #2, 4 percent was $291/ton, i.e. an average difference of $78/ton. Comparing the prices for 5 percent broken rice exported from Thailand, Vietnam and India shows that Vietnamese rice sells at a substantial discount to Thai rice, whereas the price of Indian rice is somewhat above the Thai level. Similar observations are made for 25% brokens whilst Thai A1 Super rice clearly demarks the lower bound of the low-quality range. The price series are depicted in Figures 1, 2 and 3.

Part of the explanation behind the observation that Vietnamese rice sells at a discount to Thai rice on world markets is that Vietnam entered the international rice market at a time of low world prices. Hence part of its emergence has been contingent on keeping prices competitively low, particularly in low-income markets. Even when beginning to export to higher-income markets, Vietnam has had to temper price premiums. Finally, there is the issue of low quality, which despite improved milling facilities, has been compounded by the lack of standardization systems, limited rice seed control, and insufficient drying and storage facilities. Due to internal pricing policies, India's rice is currently uncompetitive on most markets. Moreover, India has problems relating to quality and reliability much like Vietnam does. Thailand is a country with a substantially longer and more well established experience in international rice trade (especially compared to Vietnam). (See Nielsen for further discussion) There seems to be evidence that where Thailand competes with the US (on high-quality markets), it accepts a price discount. Where Thailand competes with Vietnam (on low- and medium-quality markets) it captures a price premium. Finally, it should be mentioned that apart from differences in quality, differences in transportation costs most certainly account for part of the observed price differentials.

Results

The tests of the LOP, long run exclusion and weak exogeneity are undertaken in several steps. First of all, the appropriate number of lags is determined (in a model with 11 centered seasonal dummies to take account of deterministic seasonality) starting with 12 lags (5 lags in the models that include the Indian series due to the limited number of observations) and testing this against a model with 11 lags and so forth. The conventional information criteria (Akaike and Schwarz) and the results of the sequential Likelihood Ratio (LR) tests are used to determine the appropriate lag length. Given this lag length, misspecification tests for autocorrelation, normality and autoregressive conditional heteroscedasticity (ARCH) are performed. (The results of the misspecification tests are presented in

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rather in the extent of overall integration of markets. Hence the strategy adopted here is (a) to identify the presence of seasonal unit roots in the individual data series, (b) to filter the series so that these unit roots are removed, and (c) conduct the cointegration analyses on the basis of the filtered data series. Nielsen and Yu present the results of these tests and show which filters are used on the data to deal with the identified seasonal unit roots.

[6] For the US the CPI is obtained from the US Bureau of Labor Statistics Data, for Thailand the CPI is obtained from the Bank of Thailand, for India the CPI data are obtained from the Labour Bureau (and calculated as a weighted average of the CPI for industrial workers and the CPI for agricultural and rural laborers using FAO data), and the Vietnamese CPI is obtained from the General Statistics Office.

Nielsen and Yu.) If no misspecification problems are identified, it is then tested whether or not the seasonal dummies are necessary. This is done using a LR test of the acceptability of going from a model with seasonal dummies to a model without. This test is χ^2-distributed with the degrees of freedom equal to the number of restrictions. If there are misspecification problems, however, these are corrected prior to any further testing. Autocorrelation problems are addressed by adding more lags, while normality and ARCH problems are corrected by introducing dummy variables for outliers.

Once a model has been accepted, the Trace test is used to determine the number of cointegrating relations. If the rank is found to be precisely p-1, the LOP is tested. If the rank is found to be less than p-1, tests of long run exclusion and weak exogeneity are performed. Based on the available data, the LOP is tested using a sequential nesting structure. First of all, it is tested whether there are single markets for 5% brokens and 25% brokens, respectively. To the extent that these hypotheses hold, the next step is to investigate whether or not the international market can be distinguished into a high and a low quality segment. Finally, it is tested whether there is a coherent international and inter-quality market for rice.

The market for 5% brokens

Rice with just 5% broken kernels is fairly high quality rice. As shown in Figure 2 there are substantial differences in the price levels between 5% rice exported from Thailand, Vietnam and India. Part of the reason for this difference is – as discussed above – due to perceived and real quality differences and reliability problems on the part of especially Vietnamese deliveries. Moreover, there are policy interventions that may also play an important role – particularly in the case of India (discussed by Nielsen). The next question is of course whether or not the developments of these three price series over the past years are sufficiently coherent so that they can be characterized by being driven by a common stochastic trend. Table 3 presents the results of the multivariate test of the Law of One Price.

The results of the Trace test show that there are no cointegrating relations among the three series for 5% brokens that would make the process X_t stationary. Hence the LOP cannot be tested for this group of prices. When imposing the rank of 1 onto the model and testing for long run exclusion and weak exogeneity it is found that none of the price series are excludable at the 5% level, although the Indian price series is close to being excluded. Moreover, the tests show that if the rank were 1, one would accept the hypothesis that the Indian 5% price is weakly exogenous. Although these results are weak due to the fact that the Trace test could not verify that the rank is one, these results can be interpreted as an indication of the Indian 5% series being a long run segmented market when seen in connection with the Thai and Vietnamese 5% markets.

Part of the reason why the Trace test cannot identify even just one cointegrating relation among these price series might be that the number of observations is small when we take the Indian series into account. Hence the long run variation and trends may simply not be adequately reflected in the available data. The results should therefore be interpreted with caution. In order to test whether the Thai 5% and the Vietnamese 5% series are cointegrated (and to allow for analysis based on longer time series) the next test of the LOP is performed bivariately, with the results shown in Table 4.

In this case the hypothesis of rank = 1 is accepted and the LOP can therefore be tested. The Likelihood Ratio test statistic is 5.41, which follows a a χ^2-distribution with 1 d.f., which has critical values $\chi^2_{0.95}$ =

3.84 and $\chi^2_{0.99} = 6.63$. At the usual 5% level one would therefore be inclined to reject the H_0 hypothesis that the LOP is valid, but at the 1% level it may just barely be accepted. When testing for long-run exclusion, the LR test statistics are found to be 8.30 and 14.40 for the Thai 5% and Vietnamese 5% price series, respectively. Given that these also follow the χ^2-distribution with 1 d.f., it is clear that both prices are necessary in defining the cointegration space. Moreover, the results of the weak exogeneity tests show that none of the markets seem to dominate the other. The LR test statistics are 6.22 and 7.31, respectively.

The market for 25% brokens

Rice with 25% broken kernels is of a lower quality. Once again it tested whether there is a coherent market for 25% brokens exported from Thailand, Vietnam and India. The results are shown in Table 5. As in the case of 5% brokens, the LOP cannot be tested because the Trace test can only support a rank of 1.[7] The tests for long run exclusion and weak exogeneity are shown in Table 6 for rank = 1. These results indicate that the Vietnamese 25% series is both excludable and weakly exogenous. Hence when seen in relation to the other markets for 25% brokens, the Vietnamese 25% market can be viewed as being segmented from the others.

Hence the next analysis investigates the validity of the Law of One Price on the Thai-Indian market for 25% brokens. The Trace test shows that there is one cointegrating relation between these markets and so the LOP test may be conducted. The LR test statistic for this test is 13.09, which is a clear rejection of the hypothesis (χ^2-distribution with 1 d.f.). The tests for long run exclusion and weak exogeneity lead us to conclude that none of the prices can be excluded but the Thai 25% series is on the verge of being weakly exogenous and thus we can infer that Thailand is the long run leader in this market. When testing bivariately for the markets consisting of Thai 25% and Viet 25%, and Viet 25% and Inde 25%, the Trace test strongly rejects in both cases the existence of cointegrating relations on these markets. Moreover, even if the rank of one is imposed on the models, the LOP is rejected (results are not shown here).

High-quality market

The next analysis investigates the extent to which one can speak of a single international market for high-quality rice including the following prices: US 2/4, Thai 100, Thai 5%, Viet 5% and Inde 5%. Using the Trace test to determine the rank, we find that the rank is no more than two (Table 7), and so the LOP cannot be tested. None of the variables can be excluded in the long run, but if we do impose a rank of 4 (as we would like to if we want to be able to test the LOP) then we can say that Thai 100 and Thai 5% series are weakly exogenous. In other words these series are not affected by disturbances to the equilibrium. That the Thai 100 and Thai 5% prices are non-excludable but weakly exogenous can be interpreted as these markets being long run leader markets. They contribute to the definition of the cointegrating relations, but do not respond to a deviation from the long run equilibrium experienced in the previous period. The other markets, that are both non-excludable and non-weakly exogenous, are long run follower markets in the sense that these prices play a role in determining the cointegration

[7] Imposing rank = 2 and testing for the LOP gives the following rejection: LR test = 12.59, which follows a χ^2-distribution with 2 d.f. and a resulting p-value = 0.00.

relations, but do respond to disequilibria. Estimating the model with the two Thai prices given as exogenous generates the results shown in Table 8.

The Trace test suggests that there are indeed two cointegrating relations among the variables, but the test of the LOP is rejected (the LR-test statistic being 51.22 and following a a χ^2-distribution with 6 d.f.). Hence this leads us to conclude that there are relations among the US 2/4, Viet 5%, Inde 5% markets with the Thai 100 and the Thai 5% markets being viewed as exogenous to the system, but the relations cannot be characterized by the classical Law of One Price.

Low-quality market

The first multivariate model including Thai 25%, Viet 25%, Inde 25% and Thai A1 rice does not allow us to test the validity of the LOP hypothesis because the rank is found to be just 1 (Table 9). Moreover, even if the rank of three is imposed, the LOP is strongly rejected with a LR test statistic of 26.08, following a χ^2-distribution with 3 d.f.

With the rank of one impose, the results show that the Thai 25% and the Inde 25% can be excluded from the long run relations. On the basis of these results the next test of the LOP is performed in a bivariate model including the Viet 25% series and the Thai A1 series. The results are presented in Table 10. According to the Trace test there is no stable long run relation between these two prices, but it may argued be viewing the graphs of the cointegrating relations (not shown here) that the rank might be one. Imposing this on the model and testing for the LOP actually allows us to comfortably accept the hypothesis with an LR test statistic of 1.71, which following a χ^2-distribution with 1 d.f. gives us a p-value of 0.19.

In other words, there is some (weak) evidence that the Viet 25% and the Thai A1 Super rice (which is clearly of a lower quality than Thai 25% rice) constitute one integrated market. This result seems to reaffirm the suspicion that Vietnamese rice is indeed being judged to be of a lower quality than corresponding quality grades (i.e. 25% brokens) from other countries. Tests for long run exclusion and weak exogeneity show that both series are important in determining the cointegrating relation, but that A1 is weakly exogenous. Hence, as expected due to Thailand's much longer and more stable experience on the international rice market, the A1 market is the long run leader market.

International and inter-quality market

When testing for the presence of an international and inter-quality market for rice, we cannot test the entire market because the Indian series are too short relative to the number of variables in the model. Hence the hypotheses are tested without the Indian series. Not surprisingly, the results show that we cannot test the LOP because there are not p-1 cointegrating relations. Rather there is only statistical grounding for the presence of two cointegrating relations. The exclusion tests (not shown here) reveal that the Thai 100 series and the Viet 25% series are excludable and so the model is re-estimated without these two series. Once again, the LOP cannot be tested due to a rank of just 1. The tests for weak exogeneity reveal the US 2/4 price and the Thai A1 price are weakly exogenous. In other words, they are long run leader markets. Hence the model is re-estimated with these variables exogenous.

The rank is found to be 2 (Table 11) and so the LOP may be tested. The LR test statistic is 54.09 and so with 6 d.f. in the χ^2-distribution, the hypothesis is rejected. The test for long run exclusion and weak exogeneity (Table 12) show that all the included series are important in defining the long run relations, and that none of them are weakly exogenous. Hence they all seem to be long run follower markets. Hence, on the basis of the results of the model estimations it seems that we can conclude that the US 2/4 price is a long run leader in the high quality market while the Thai A1 price is the long run leader in the low quality market.

Finally, because of the evidence that the development of the price of Vietnamese 25% rice could actually be seen to be more closely related to the development of the price of Thai A1 rice rather than Thai 5% rice, a bivarate model is estimated to see whether this relation holds in the higher quality range, namely between Viet 5% and Thai 25%. The results are shown in Table 13. The rank is indeed found to be one so the LOP can be tested, but with a LR test statistic of 8.64 it is rejected when held up against a χ^2-distribution with 1 d.f. Performing long run exclusion and weak exogeneity tests, we find that the Thai 25% is weakly exogenous (LR test statistic 0.21 which follows a χ^2-distribution with 1 d.f.) and hence we can conclude, as expected that it is the Thai price that takes the long run leader role on this market.

Preliminary conclusions

The results of the co-integration analysis strongly suggest that the international rice market is highly segmented: the Law of One Price was rejected strongly in most cases. In the 5% market there was weak evidence that the Thai and Vietnamese markets could be integrated although the test results are sensitive to the level at which the test is performed. More interestingly, the results confirm two of our suppositions. The first is that the Thai market dominates the international rice market. In many cases the cointegration results confirm that the Thai market takes on a long run leader role – a clear reflection of its many years of being the world's leading rice exporter. The second supposition to be confirmed is that Vietnamese rice is considered to be of a lower quality when being compared with rice of a similar grading from other countries. The results suggest that the market for Vietnamese 25% rice is not integrated with e.g. the Thai 25% market, but rather the lower quality Thai A1 rice market. Notwithstanding the weak evidence of the few exceptions mentioned above, it is safe to conclude that the Armington assumption of heterogenous goods is valid for the international rice market. Moreover, the results suggest that these products are imperfect substitutes and hence the next step of the analysis is to estimate the elasticities of substitution among these different sources.

Estimating Elasticities Of Substitution in the International Rice Market

As mentioned earlier the Armington approach is often used in global CGE models, including e.g. the standard GTAP model. Under the Armington approach, each domestically produced good is modeled as an imperfect substitute for the corresponding composite imported good. Furthermore, imported goods from different sources are modeled as imperfect substitutes for one another. Typically, these two types of imperfect substitutability are specified in Constant Elasticities of Substitution (CES) functions. The size of the substitution parameter in the CES function has a significant impact on the changes in magnitude and direction of trade flows when a policy shock is simulated. In the existing empirical literature, a few studies have estimated the substitutability between composite imports and domestic goods for a variety of countries and goods (see e.g. Stern, Francis and Schumacher; Shiells, Stern and

Deardorff; Shiells and Reinert; Reinert and Roland-Holst; Kapuscinski and Warr; and Gallaway, McDaniel and Rivera). However, empirical studies on the substitutability between imports from different sources are rare. Given the importance of the elasticities of substitution and the fact that these parameters have only rarely been estimated, we push our efforts one step further—we directly estimate the elasticities of substitution between sources of imports. By quantifying the degree of differentiation of the rice market, these estimates will provide the opportunity of better assessing the effects of trade liberalization in the rice market.

Data

One possible reason for the relatively little attention devoted to estimate elasticites of substitution between import sources is the lack of times series data with sufficient commodity and trading partner details. Indeed, it is very difficult to assemble a long enough data set with a consistent list of exporters. Nevertheless, we have been able to collect monthly import values (c.i.f.), quantities and unit values of rice trade for several major importers for two HS 6-digit commodities: semi- or wholly-milled rice (HS100630) and broken rice (HS100640).[8] These two 6-digit items comprise the processed rice sector in the GTAP database. Hence, in our estimation, we also aggregate them into a single good. These data are available for the European Union (EU), the USA, Japan, Brazil and Indonesia. Their respective groups of trading partners are very diverse. The length of the series ranges from 85 monthly observations for Japan to 60 observations for Brazil. A summary of these series is listed in Table 14.

The most disturbing feature of the data set is that there are many small exporters that sporadically enter the market. The many observations with zero value make it difficult to estimate the substitutability among these suppliers. As such, our focus is on those exporters that have significant trade shares and that are stable suppliers, while the others are combined and treated as a single region (Rest of World, ROW) in the estimation. The major exporters are identified for each importer (see Table 14). For example, India, Pakistan, the USA and Thailand are the most stable suppliers for the EU, while for Brazil, it is Argentina, Uruguay and the USA that are the major trade partners. Even the major exporters did not export rice to the corresponding importers every month. Although not surprising, given the fact that the world rice market is thin and is essentially a residual market, this certainly poses a problem for our estimation. We are forced to discard the observations with zero values and end up with fewer observations in the estimation.

Estimation Methodology

We start the discussion with the definition of a pair-wise elasticity of substitution between imports from sources *i* and *j*. Let q_i, q_j, p_i and p_j be the quantities and prices for imports from sources *i* and *j*, respectively. The elasticity of substitution σ_{ij} can be defined as:

$$\sigma_{ij} = \frac{\partial \ln(q_i / q_j)}{\partial \ln(p_j / p_i)} \qquad (3)$$

It measures how the quantity ratio of imports from *i* over imports from *j* adjusts when there is a marginal change in the price ratio of imports from *j* over imports from *i*. Thus a large positive value of

[8] The data have been extracted from the World Trade Atlas database.

σ_{ij} indicates a high degree of flexibility in substituting imports from source i with imports from source j, and vice versa. To operationalize (3), an econometric model can be specified as:

$$Q_{ij}(t) = a_{ij} + b_{ij} P_{ij}(t) + u_{ij}(t) \tag{4}$$

where $Q_{ij(t)}$ is the log of the ratio of quantity i over quantity j for period t, and $P_{ij(t)}$ is the log of the ratio of price j over price i for period t. The error term is $u_{ij(t)}$ is and a_{ij} is a constant. Coefficient B_{ij} is the elasticity of substitution between imports from sources i and j. For an exporter with n trading partners, there are $(n-1)^2$ elasticities of substitution to be estimated. However, by the construction of equation (4), substitution between sources i and j is the same as the substitution between j and i. Thus, there are only $(n-1)^2/2$ pair-wise elasticities to be estimated. As such, we only estimate model (4) for those pairs with $i<j$, while the cases for $i>j$ can be derived from symmetry.

Ordinary Least Square (OLS) estimation of (4) will generate spurious estimates when the time series used are non-stationary. As discussed earlier in this paper, the price series used in our co-integration analysis are tested for stationarity and indeed the results confirm the integration of order one of these series, I(1). In this section, similar tests are applied to the price and quantity series, as well as to the logarithms of the price and quantities ratios. Again, most of the series are shown to be I(1). The model is therefore reformulated as an Error Correction Model (ECM, proposed by Engle and Granger), i.e.:

$$\Delta Q_{ij}(t) = a_{ij} + \gamma_{ij}^{1} \Delta P_{ij}(t) + \gamma_{ij}^{2} Q_{ij}(t-1) + \gamma_{ij}^{3} P_{ij}(t-1) + u_{ij}(t) \tag{5}$$

where Δ is the difference operator. Coefficient γ^{1}_{ij} is then the short run elasticity, while the long run elasticity can be computed as $-\gamma^{3}_{ij} / \gamma^{2}_{ij}$.

Estimation Results

Estimation of equation (5) generates both short run and long run elasticities. The results are displayed in Table 15. Except for Japan, most of the short-run elasticities are significant. Specifically, the estimates for the USA, Brazil and Indonesia are significant at the 1 per cent level. Values for most of the significant short-run elasticities are in the range of 0.7 to 2, which are much smaller than the ones used in many CGE models (e.g. in GTAP, these are around 4) and indicate very low substitutability among exporters. The estimates of short-run elasticities for Japan as an importer are mostly insignificant. In fact, 4 of the 6 pairs show negative signs. This is possibly due to many zero observations in the Japanese data set and the fact that only less than half of the observations are used in the estimation (34 observations out of a total of 85). As such, we opt to not use the estimates for Japan in the subsequent illustrative examples.

Compared to the relatively robust estimates of the short-run elasticities, estimates of the long-run elasticities are less satisfactory. Many of the long run elasticities have negative signs and the corresponding estimated parameters are not significant. This is especially true for the EU and Japan—there are only two cases that are significant at the 1 per cent level for the EU, while none of the estimates for Japan are significant. Where the estimates are significant, the long-run elasticities are not very different from the short-run elasticities, i.e. they are either of the same magnitude as the short-run ones or slightly larger. In a few cases, the long-run elasticities are smaller. The length of the time series used in this estimation may explain the failure in identifying satisfactory long-term relationship. On the other hand, the complexity of the rice market and the thinness of rice trade may also make this difficult. The simple specification of our model certainly does not take in account any other possible explanatory variables related to e.g. policy instruments and consumer preferences. Due to the statistical properties of the long-run elasticities and the fact that they are not very different from the short-run elasticities, which in generally are robust, we choose to apply the short run elasticities in our subsequent illustrative CGE simulations.

Implications for trade liberalization

Trade liberalization in the form of tariff removal will reduce the border price and thus attract more imported goods. The substitution elasticities between composite imports and domestically produced goods (denoted as *ESUBD* hereafter) will determine how the composite imported good substitutes for the domestically produced good, while the substitution elasticities between import sources (denoted as *ESUBM* hereafter, which is the one estimated here) will determine how well imports from different sources substitutes one another. Given the strong evidence of low short-run elasticities of substitution, one would expect the adjustment of trade shares of rice from different sources to be rigid when the price of imported rice decreases. We will examine this hypothesis in this section using a CGE model.

Illustrative examples

To illustrate the importance of using the correct size of these parameters, we conduct a simple trade liberalization scenario in which all the regions completely remove their import tariffs on processed rice and paddy rice. This experiment is simulated with three different sets of parameters (see Box 1), using the standard GTAP model. The first set of parameters is drawn from the standard GTAP parameter file, with ESUBM being 4.4 for rice. The second set of the parameters comes from the estimation in this study. The CES functional forms are used for both Armington nests in the GTAP model, which implies that there is only one substitution parameter for each importer. Thus, we apply from our estimation the average[9] ESUBM (across all the pair-wise estimates) for each importer. For importers that are not included in our estimation, the average across the four countries in the estimation (Japan excluded) is used. For commodities other than rice, the default ESUBMs from GTAP are used. The last set of parameters contains the same ESUBMs as in the second set. In addition, the ESUBDs for rice are reduced to half of the average of the ESUBMs[10]. This is intended to illustrate the different effects of ESUBD and ESUBM in determining the volume and direction of rice trade. It also complements our analysis since we do not offer any estimation for ESUBD.

Simulation results

First, we look at the percentage changes in total processed rice trade by importers under the three experiments. These are shown in Table 16. As expected, in all three experiments most of the regions

[9] By applying the averages of the pair-wise elasticities of substitution for each importer, invaluable information from our estimation is lost. We are forced to do this because of the CES characterization of the Armington structure in the CGE model that we choose for our simulations. The non-nested CES specification is parsimonious since it only has one substitution parameter. While this is a practical choice for large-scale multi-sector and multi-national models, for a study such as ours, nested CES or other functional forms (such as the Almost Ideal Demand System) that allows for different elasticities will produce more sensitive results by taking advantage of our pair-wise estimates.

[10] This relationship between ESUBD and ESUBM is supported by some empirical studies. See e.g. Jomini et al. and Liu, Arndt and Hertel.

increase their rice imports as a result of the tariff cut. The most notable increase occurs in Japan, where the initial tariff rate listed in the GTAP database is more than 400 percent. The few exceptions are Indonesia and the USA, where imports decease slightly due to the very low initial tariff rates compared with other rice importers. The universal removal of tariffs therefore results in some trade diversion on the part of these two countries.

With the same ESUBD in both Exp1 and Exp2, the substitutability of imported rice for domestic rice remains the same - thus change in total imports should be more or less of similar magnitudes in these two cases. Comparing Exp1 and Exp2, however, one can see that the resulting increases in total imports in Exp2 are smaller than in Exp1 for most regions (for Indonesia and USA, the decrease is greater). This is because the rigidity introduced in Exp2 by means of smaller values of ESUBM makes it difficult to substitute between sources and this has a non-negligible impact on total imports. A comparison of Exp1 and Exp2 against Exp3, where ESUBD is reduced to half of the size of ESUBM, shows that indeed ESUBD is more crucial in determining the size of total imports. Again, the most notable example is Japan, whose total imports increase by only 149 percent, in comparison to over 1300 percent in the first and second experiment. This certainly calls for more empirical studies on the size of ESUBD.

We now turn our attention to the trade shares by source for each importer. In Table 17, the initial trade shares from the base data and the post-simulation shares are listed for several regions. The persistence of trade shares is most evident for the major trading partners for each importer. For Japan, the major partners are Thailand and USA (with trade shares of 0.248 and 0.340 respectively). Under Exp1, Thailand’s share falls to 0.191 while USA’s share increases to 0.382. Under Exp2, however, changes in trade shares for Thailand and USA are very small—for Thailand it is 0.245 and for USA it is 0.341. Similar results are found for other importers. For example, for the USA, Thailand’s share falls from 0.525 to 0.430 under Exp1, while under Exp2 and Exp3, Thailand’s shares are, respectively, 0.516 and 0.523. For Indonesia, Thailand’s share declines from 0.475 to 0.399 under Exp1, while under Exp2 and Exp3, Thailand’s shares are, respectively, 0.458 and 0.471. For the EU, adjustment of trade shares is generally larger than for other importers. This is particularly so for the share of intra-EU trade, which falls from 0.608 to 0.116 under Exp1 and becomes 0.397 and 0.394 under Exp2 and Exp3, respectively. The removal of EU’s tariff certainly contributes to this large switch from internal trade to external sources. Nonetheless, this result also confirms the general observation of much more persistent trade shares when applying smaller values of ESUBM.

Overall, we observe that trade shares are more persistent when the estimated values of ESUBM are applied, i.e. under Exp2 and Exp3. A further comparison of Exp2 and Exp3 reveals hardly any difference in trade shares, despite very different changes in total imports under these two cases. This suggests that trade shares – as expected – are much more sensitive to the size of ESUBM that that of ESUD. To summarize these observations, an index similar to the mean square error measure is computed to measure the deviation of trade shares resulting from the three experiments from the base data. This is reported in the last row of Table 17. This index is substantial under Exp1 for all the five importers, ranging from 50.89 for the EU, to 10.19 for the USA, to 0.93 for Brazil. Under Exp2 and Exp3, it falls to below 1 for Japan (0.39 and 0.08), USA (0.97 and 0.22) and Brazil (0.04 and 0.02). For the EU, it remains over 20 in both Exp2 and Exp3, mainly due to the switch from intra-EU to extra-EU trade.

Concluding remarks

When evaluating the impact of trade policy reforms using e.g. a computable general equilibrium model, it is clear that the underlying data, including the parameters, are of key importance for the results. One of the critiques that has been raised against some CGE models is the lack of econometric foundation for the behavioral parameters used in the models. This paper has attempted to contribute to the continued efforts to supply CGE models with a more solid econometric foundation. Our efforts focus on one type of elasticity – the Armington elasticity of substitution among imports from different sources, and one commodity – rice. Before embarking on the estimation of the elasticities, a cointegration analysis has been used to determine whether and to what extent the Armington assumption of differentiated products is relevant for the international rice market. The results clearly support this assumption in that it is found (with a few border-case exceptions) that the Law of One Price – by which products could be viewed as homogenous – is strongly rejected in virtually all segments of the international rice market. Two suppositions, however, were confirmed by the cointegration analysis: (a) the Thai market dominates the long-run price formation process in several quality segments of the market, and (b) Vietnamese rice is being judged as being of a lower quality than similar grade rice from other countries.

Time series rice trade data is then applied to an Error Correction Model to estimate the substitutability between sources of imports for several major importers. Our estimation results reveal much lower values than the ones frequently used in CGE models. This finding is consistent with the results of the cointegration analysis, which suggests that the international rice market is highly segregated and therefore importing countries treat rice from different sources as differentiated products. Hence it is concluded that the Armington trade structure frequently applied in CGE models is highly appropriate for the rice sector but that one should reconsider the size of the elasticities of substitution currently being used. Applying the – lower – econometrically estimated elasticities to trade liberalization experiments using a fairly standard global CGE model generates much more persistent trade structures (in terms of import shares) than when the same experiments are performed using the – higher – conventional elasticities. Furthermore, our examples also show that while the elasticity of substitution between domestically produced rice and the composite imported rice has a decisive role in determining the change in total rice imports, it plays no significant role in determining the shares of imports among sources.

Despite the strong conclusions regarding the acceptance of the Armington approach, the size of the elasticity of substitution among different sources for imports, and the general implications for trade liberalization, several limitations remain and hence point to further research. First, the data series used in both the cointegration analysis, but particularly also in the estimation of elasticities, are not very long, and so the long-term relations may not be adequately captured. In the estimation of elasticities, we could therefore not generate a complete set of long-run elasticities. Second, it is acknowledged that the international rice market is characterized by a wide range of protective policy instruments and strong consumer preferences that might be considered included in the model. However, to compile such information over a longer period of time is not a trivial issue. Third, although seasonality is addressed in the cointegration analysis, it is not addressed in the estimation of the elasticities. This is clearly an issue that needs to be tackled. Finally, when applying the estimated elasticities to the CGE model, the average is used due to the restrictive nature of the CES functional form in the model, despite the fact that our pair-wise estimates yield different elasticities. Such information would no doubt affect the simulation results. Ideally, alternative functional forms that can be calibrated to differential pair-wise elasticities should be used. Despite these caveats, this analysis has rather robustly confirmed the application in CGE models of the Armington approach to describing international rice trade. Furthermore, it has made clear that the elasticities of substitution between imports from different sources should be substantially lower than the ones currently being applied in many standard CGE models.

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Tables

TABLE 1. **Classification of markets according to results of exclusion and exogeneity tests**

	Long run exclusion	**Weak exogeneity**	**Interpretation**
Long run segmented market	√	√	Price series that are excluded from long run relations and also do not respond to disequilibria.
Long run leader market	×	√	Price series that contribute to the definition of the cointegrating relations, but do not respond to deviations from the long run equilibrium.
Long run follower market	×	×	Price series that contribute to the definition of the long run cointegrating relations and also are responsive to disequilibria.
Long run regulator market	√	×	Price series that are excluded from long run relations but for some reason do respond to disequilibria.

Note: A √ denotes an acceptance of the hypothesis whilst an × denotes a rejection. Source: Adapted from Le Goulven (1999)

TABLE 2. **Summary of data used in the cointegration analysis**

	US 2/4	Thai 100	Thai 5	Viet 5	India 5	Thai 25	Viet 25	India 25	Thai A1
Series length	1990:01 2001:12	1990:01 2001:12	1990:01 2001:12	1990:01 2001:12	1996:01 2001:12	1990:01 2001:12	1990:01 2001:12	1996:01 2001:12	1990:01 2001:12
Observations	144	144	144	144	72	144	144	72	144
Average (real) price USD/ton	291	213	205	112	220	167	90	198	86

TABLE 3. **Multivariate test of LOP for 5% brokens: Thailand, Vietnam and India**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 96:11-01:12	21.23	p = 0	3	31.76	29.38	26.70
Lags = 5, 11 seasonal dummies	3.75	p ≤ 1	2	17.24	15.34	13.31
1 dummy for outliers, 62 obs.	0.15	p ≤ 2	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 4. **Bivariate test of LOP for 5% brokens: Thailand and Vietnam**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 90:09-01:12	16.53**	p = 0	2	17.24	15.34	13.31
Lags = 5, 11 seasonal dummies						
1 dummy for outliers, 136 obs.	0.36	p ≤ 1	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 5. **Multivariate test of LOP for 25% brokens: Thailand, Vietnam and India**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 96:07-01:12	46.17***	p = 0	3	31.76	29.38	26.70
Lags = 3, 11 seasonal dummies	11.92	p ≤ 1	2	17.24	15.34	13.31
1 dummy for outliers, 66 obs.	0.00	p ≤ 2	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 6. **Tests of long run exclusion and weak exogeneity for 25% brokens: Thailand, Vietnam & India**

Variable	Long run exclusion	Weak exogeneity	d.f.	χ^2 critical value
Thai 25%	9.38**	5.06**	1	3.84
Viet 25%	0.37	0.19		
Inde 25%	21.85**	17.84**		

Note: ** indicates significance at the 5 percent level.

TABLE 7. **Multivariate test of LOP for high quality rice: US 2/4, Thai 100, Thai 5%, Viet 5%, Inde 5%**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 96:11-01:12	123.07***	p = 0	5	72.21	68.68	64.74
Lags = 5, 11 seasonal dummies	63.81***	p ≤ 1	4	50.19	47.21	43.84
No dummies, 62 observations	21.95	p ≤ 2	3	31.76	29.38	26.70
	5.17	p ≤ 3	2	17.24	15.34	13.31
	1.76	p ≤ 4	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 8. **Multivariate test of LOP for high quality rice:**
US 2/4, Viet 5%, Inde 5%: endogenous series; Thai 100, Thai 5%: exogenous series

Model	Trace test	H_0: Rank = p	P – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 96:11-01:12	71.84***	p = 0	3	31.76	29.38	26.70
Lags = 5, 11 seasonal dummies	20.27***	p ≤ 1	2	17.24	15.34	13.31
No dummies, 62 observations	4.27**	p ≤ 2	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 9. **Multivariate test of LOP for low quality rice: Thai 25%, Viet 25%, Inde 25%, Thai A1**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 96:09-01:12	54.10***	p = 0	4	50.19	47.21	43.84
Lags = 5, 11 seasonal dummies	19.65	p ≤ 1	3	31.76	29.38	26.70
1 dummy for outliers, 64 obs.	5.69	p ≤ 2	2	17.24	15.34	13.31
	0.03	p ≤ 3	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 10. **Bivariate test of LOP for low quality rice: Vietnam 25% and Thai A1**

Model	Trace test	H_0: Rank = p	p – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 90:10-01:12 Lags = 6, No seasonal dummies	10.92	p = 0	2	17.24	15.34	13.31
1 dummy for outliers, 135 obs.	2.07	p ≤ 1	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 11. **Multivariate test of LOP for international and inter-quality rice:**
Thai 5%,Viet 5%,Thai 25% endogenous; US 2/4 and Thai A1 exogenous

Model	Trace test	H_0: Rank = p	P – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 90:09-01:12	98.23***	p = 0	3	31.76	29.38	26.70
Lags = 5, 11 seasonal dummies	17.60***	p ≤ 1	2	17.24	15.34	13.31
1 dummy for outliers, 136 obs.	3.06	p ≤ 2	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 12. **Tests of long run exclusion and weak exogeneity for model with Thai 5%,Viet 5%,Thai 25% endogenous; US 2/4 and Thai A1 exogenous**

Variable	Long run exclusion	Weak exogeneity	d.f.	χ^2 critical value
Thai 25%	74.32**	18.16**		
Viet 25%	54.18**	21.89**	2	
Thai A1	73.73**	26.38**		5.99
US 2/4	23.09**	-		
Thai A1	41.33**	-		

Note: ** indicates significance at the 5 percent level.

TABLE 13. **Bivariate test of LOP for V5 and T25**

Model	Trace test	H_0: Rank = p	P – r	Critical value 99%	Critical value 95%	Critical value 90%
Effective sample: 90:09-01:12 Lags = 5, 11 seasonal dummies	23.10***	p = 0	2	17.24	15.34	13.31
1 dummy for outliers, 136 obs.	1.02	p ≤ 1	1	5.02	3.84	2.71

Note: ***/**/* indicates significance at the 1,5 and 10 percent levels, respectively.

TABLE 14. **Summary of the data used for the estimation**

Importers	EU	USA	Japan	Brazil	Indonesia
Major partners	India, Pakistan, USA, Thailand	India, EU, Pakistan, Thailand, China	China,Thailand, USA	Argentina, Uraguay, USA	India, Pakistan, Thailand, Vietnam
Series length	1996:1-2001:7	1995:1-2001:12	1994:2-2001:2	1997:1-2001:12	1996:1-2001:10
Observations*	67	84	85	60	70

Source: the World Trade Atlas.
*: Due to numerous holes in the data set, the actual number of observations used is less than the one listed here, especially for Japan and Indonesia.

TABLE 15. **Estimated elasticities of substitution among importers**

Importer	Exporters	Short-run	Long-run	R Square
EU	India-Pakistan	1.167 ***	0.673	0.474
	India-Thailand	1.031 ***	0.724	0.465
	India-USA	1.807 ***	**1.987** ***	0.443
	India-ROW	0.397	-0.543	0.518
	Pakistan-Thailand	0.572 *	-0.211	0.449
	Pakistan-USA	1.699 ***	1.127 **	0.571
	Pakistan-ROW	1.088 **	-1.219	0.477
	Thailand-USA	2.203 ***	1.999 ***	0.472
	Thailand-ROW	0.974 *	-1.970 *	0.347
	USA-ROW	3.336 ***	2.141	0.446
USA	India-EU	0.835 ***	0.675 ***	0.611
	India-Pakistan	0.962 ***	0.921 **	0.614
	India-Thailand	0.962 ***	0.868 **	0.727
	India-China	1.650 ***	1.526 ***	0.733
	India-ROW	1.362 ***	**1.401** **	0.612
	EU-Pakistan	0.886 ***	0.726 *	0.519
	EU-Thailand	0.692 ***	0.472	0.446
	EU_China	1.997 ***	**2.055** ***	0.647
	EU-ROW	1.622 ***	**1.874** ***	0.608
	Pakistan-Thailand	0.905 ***	0.670	0.513
	Pakistan-China	2.110 ***	**2.330** ***	0.678
	Pakistan-ROW	1.800 ***	**2.134** ***	0.598
	Thailand-China	1.925 ***	1.922 ***	0.680
	Thailand-ROW	1.814 ***	**2.306** ***	0.649
	China-ROW	1.823 ***	1.774 ***	0.599
Japan	China-Thailand	-0.257	-0.657	0.368
	China-USA	1.852 *	1.588	0.517
	China-ROW	-0.731	-0.047	0.483
	Thailand-USA	0.578	0.265	0.601
	Thailand-Row	-0.604	-0.550	0.528
	USA-ROW	-0.772	-0.650	0.485

Source: Estimation results.
Note: ***, ** and * denote significance at 1, 5 and 10 per cent, respectively. The numbers in bold in the fourth column signal that these are larger than the corresponding short-run elasticities.

TABLE 15 (cont.). **Estimated elasticities of substitution among importers**

Importer	Exporters	Short-run	Long-run	R Square
Brazil	Argentina-Uruguay	1.026 ***	0.919	0.331
	Argentina-USA	0.874 ***	0.874 **	0.700
	Argentina-Row	2.177 ***	1.803 ***	0.782
	Uruguay-USA	0.920 ***	**1.018** ***	0.692
	Uruguay-ROW	2.183 ***	2.003 ***	0.831
	USA-ROW	1.596 ***	**2.005** ***	0.705
Indonesia	Thailand-Vietnam	1.689 ***	1.623 ***	0.627
	Thailand-ROW	1.574 ***	**2.964** **	0.504
	Vietnam-ROW	1.996 ***	**2.427** ***	0.581

Source: Estimation results.
Note: ***, ** and * denote significance at 1, 5 and 10 per cent, respectively. The numbers in bold in the fourth column signal that these are larger than the corresponding short-run elasticities.

TABLE 16. **Percentage changes in total imports of rice, from the base data**

	AUS	CHN	JPN	IDN	THA	VNM	REA	IND	XSA	USA	EU	ARG	BRA	URY	MEA	XRW
Exp1	3.9	69.3	1412	-11.6	34.0	36.0	49.6	3.7	63.6	-0.2	59.1	20.4	31.4	33.4	14.7	33.8
Exp2	1.0	54.8	1380	-15.1	45.1	30.0	41.0	5.1	63.5	-4.1	44.8	20.1	31.3	33.0	13.8	32.3
Exp3	0.2	20.0	149.0	-1.1	2.3	5.8	15.0	0.7	14.7	1.9	14.2	5.0	8.5	7.9	4.3	8.8

Sources: Simulation results.

Box 1
Illustrative trade liberalization scenario with 3 different sets of parameters: 100% import tariff cut for processed and paddy rice in all regions

Experiments

Experiment 1	Default GTAP parameters (ESUBD and ESUBM)
Experiment 2	Averages of estimated ESUBM for USA, EU, Indonesia and Brazil are used for these regions for both processed and paddy rice; averages across the above four regions are used for all the other regions; ESUBD set to GTAP default value
Experiment 3	Same as Experiment 2 except ESUBD for processed and paddy rice set to half of the ESUBM.

Changes to the GTAP model

ESUBM(i) changed to ESUBM(i,s) where *i* is the index for trade commodities and *s* is the destination. In other words, ESUBM is now destination specific. Recall that the averages of the pair-wise ESUBMs estimated in this study are used for each *s*.

Regional and Commodity Aggregation of the GTAP database

16 regions/countries	Australia, China, Japan, India, Thailand, Vietnam, Rest of East Asia, Indonesia, Rest of South Asia, USA, EU, Argentina, Brazil, Uruguay, Mid East and North Africa, and the rest of the world.
8 goods	Paddy and processed rice, other crops, livestock and meat products, other food, other primaries, manufactures, and services.

TABLE 17. **Import shares by sources for selected rice importers under alternative elasticities of substitution**

	Japan				USA				EU				Brazil				Idn.			
	Base	Exp1	Exp2	Exp3	Base	Exp1	Exp2	Exp3	Base	Exp1	Exp2	Exp3	Base	Exp1	Exp2	Exp3	Base	Exp1	Exp2	Exp3
AUS	0.188	0.173	0.187	0.188	0.000	0.000	0.000	0.000	0.007	0.013	0.010	0.010	0.000	0.000	0.000	0.000	0.049	0.049	0.049	0.049
CHN	0.059	0.067	0.060	0.059	0.040	0.048	0.040	0.040	0.031	0.075	0.048	0.048	0.004	0.004	0.004	0.004	0.054	0.067	0.056	0.055
JPN	0	0	0	0	0.022	0.032	0.023	0.023	0.017	0.049	0.026	0.026	0.002	0.003	0.002	0.002	0.004	0.006	0.004	0.004
JDN	0.052	0.057	0.052	0.052	0.046	0.053	0.047	0.046	0.034	0.081	0.053	0.053	0.005	0.005	0.005	0.005	0	0	0	0
THA	0.248	0.191	0.245	0.248	0.525	0.430	0.516	0.523	0.091	0.150	0.134	0.139	0.016	0.011	0.017	0.016	0.475	0.399	0.458	0.471
VNM	0.001	0.001	0.001	0.001	0.071	0.070	0.071	0.071	0.001	0.002	0.001	0.001	0.031	0.026	0.031	0.031	0.150	0.152	0.153	0.150
REA	0.031	0.036	0.032	0.031	0.030	0.037	0.031	0.030	0.020	0.050	0.032	0.031	0.003	0.004	0.003	0.003	0.006	0.007	0.006	0.006
IND	0.005	0.006	0.005	0.005	0.125	0.148	0.127	0.125	0.017	0.040	0.026	0.026	0.000	0.000	0.000	0.000	0.001	0.001	0.001	0.001
XSA	0.005	0.005	0.005	0.005	0.026	0.031	0.026	0.026	0.007	0.016	0.011	0.011	0.000	0.000	0.000	0.000	0.249	0.303	0.259	0.252
USA	0.340	0.382	0.342	0.341	0	0	0	0	0.058	0.139	0.089	0.089	0.009	0.010	0.009	0.009	0.001	0.002	0.001	0.001
EU	0.006	0.009	0.006	0.006	0.048	0.070	0.049	0.048	0.608	0.116	0.397	0.394	0.004	0.004	0.003	0.003	0.001	0.002	0.001	0.001
ARG	0.000	0.000	0.000	0.000	0.004	0.004	0.004	0.004	0.000	0.000	0.000	0.000	0.417	0.422	0.416	0.416	0.000	0.000	0.000	0.000
BRA	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.002	0.001	0.001	0	0	0	0	0.000	0.000	0.000	0.000
URY	0.001	0.001	0.001	0.001	0.000	0.000	0.000	0.000	0.001	0.003	0.002	0.002	0.497	0.499	0.497	0.497	0.000	0.000	0.000	0.000
MEA	0.004	0.004	0.004	0.004	0.004	0.005	0.004	0.004	0.003	0.007	0.005	0.005	0.001	0.001	0.000	0.000	0.001	0.001	0.001	0.001
XRW	0.059	0.067	0.059	0.059	0.059	0.071	0.060	0.060	0.106	0.258	0.165	0.164	0.011	0.011	0.011	0.011	0.010	0.012	0.010	0.010
		7.30	0.39	0.08		10.19	0.97	0.22		50.89	21.99	22.33		0.93	0.04	0.02		9.38	2.05	0.43

Source: GTAP database and simulation results.

Note: Base refers to the import shares by sources from the GTAP database. Exp1, Exp2 and Exp3 refer to, respectively, the post-simulation shares from the three simulations. The last row measures the deviation of import shares under the three experiments from the base case ones. These are calculated as

$$\left\{\sum_r \left[(s^i_{rs} - s^b_{rs})*100\right]^2\right\}^{½}$$

, where S_{rs} denotes the share of import of destination s from source r, *i* denotes the *i*th experiment and *b* denotes the base case.

Figures

FIGURE 1. **CPI-deflated prices for US 2/4 and Thai 100 rice, US$/ton, 1990-2001**

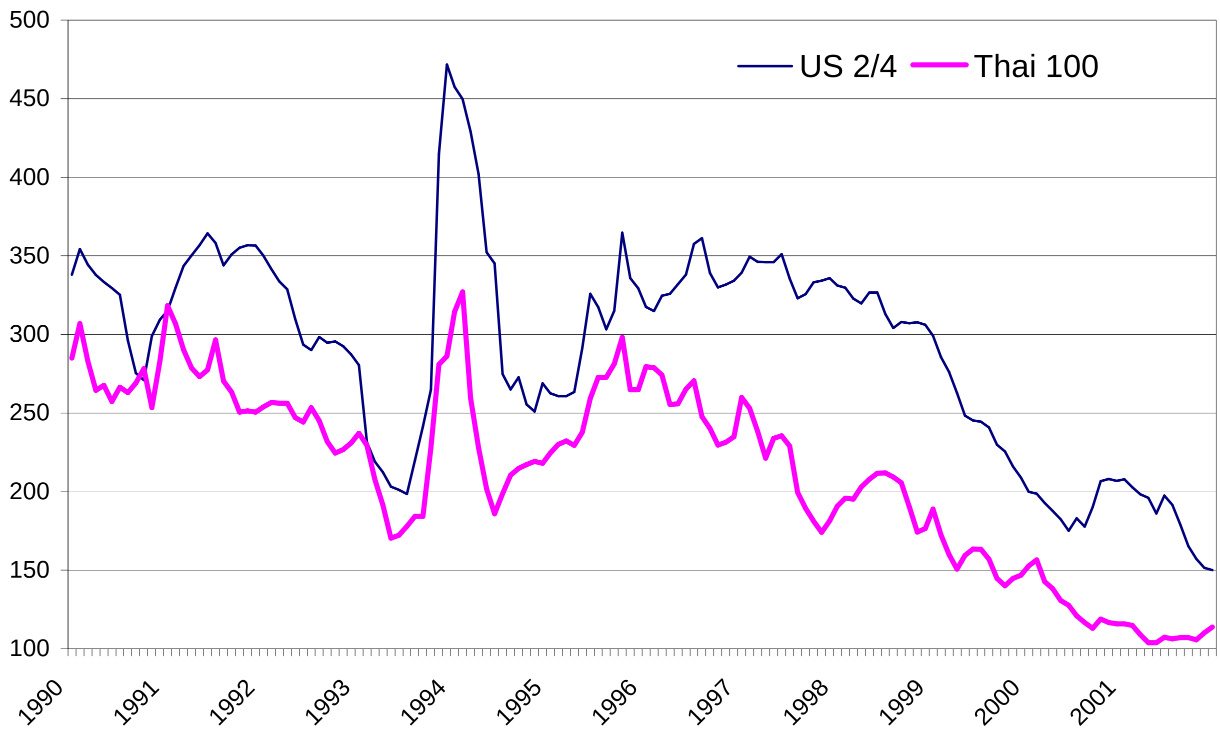

FIGURE 2. **CPI-deflated prices for 5% broken rice, US$/ton, 1990-2001**

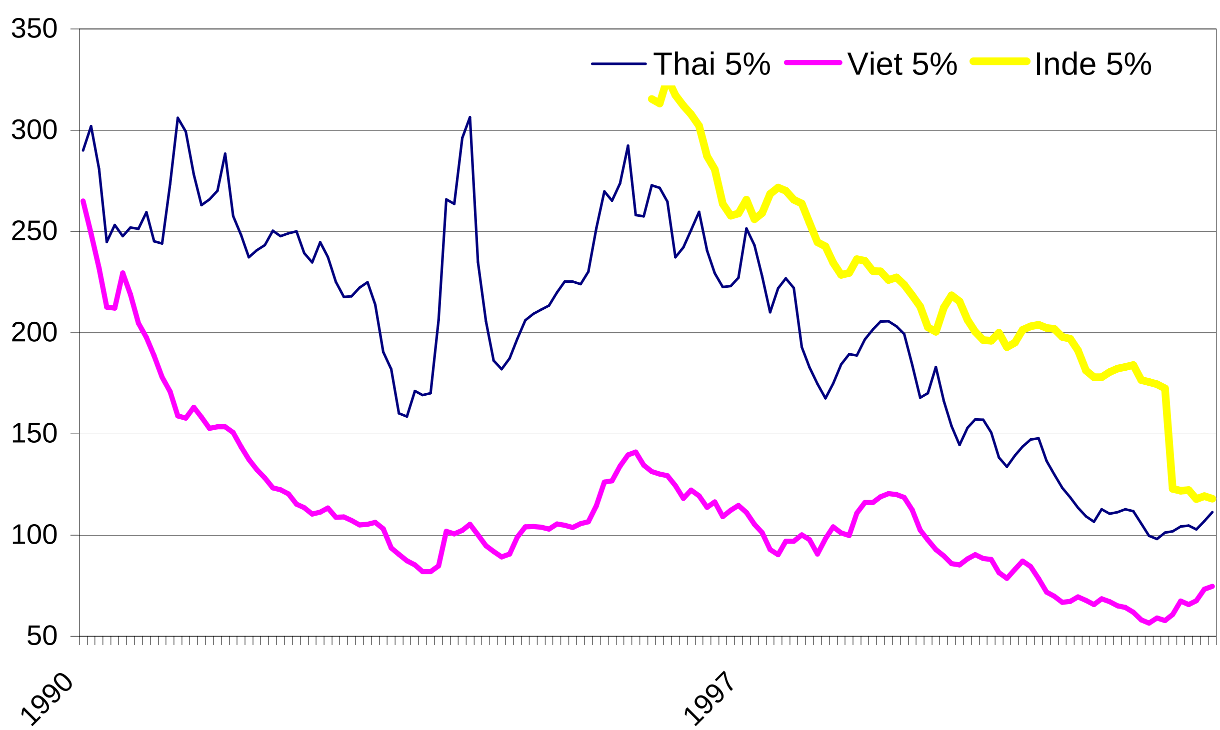

FIGURE 3. **CPI-deflated prices for 25% broken rice and Thai A1 rice, US$/ton, 1990-2001**