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Measuring market power in the French Comté cheese market

Pierre R. Mérel

University of California, Davis, CA, USA

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Abstract

A new empirical industrial organisation approach is used to measure seller market power in the French Comté cheese market, characterised by government-approved supply control. The estimation is performed on quarterly data at the wholesale stage over the period 1985–2005. Three different elasticity shifters are included in the demand specification, and the supply equation accounts for the existence of the European dairy quota policy. The market power estimate is small and statistically insignificant. Monopoly is clearly rejected. Results appear to be robust to the choice of functional form and suggest little effect of the supply control scheme on consumer prices.

Keywords: supply control, market power, NEIO, Comté cheese, protected designation of origin

JEL classification: L13, L41, Q18

1. Introduction

Comté is one of the most popular cheeses in France, with an annual production of about 50 thousand metric tons, making the Comté industry the largest cheese industry benefiting from a protected designation of origin (PDO) in the country.[1] Production has been increasing steadily since the early 1990s. The industry is characterised by the existence of industry-wide contracts between upstream producers (dairy cooperatives) and downstream processors (ripening facilities), as well as government-approved supply control.

In this article, a new empirical industrial organisation (NEIO) approach is used to assess the degree of seller market power exercised in the Comté cheese market. The question has policy relevance given the importance of the industry and the controversial nature of the supply control scheme. In 1998, the producer association, the *Comité Interprofessionnel du Gruyère de Comté* (hereafter CIGC), was fined by the French antitrust authority, the *Conseil*

*Review coordinated by Alison Burrell.

1 Most of the production is consumed nationally.

de la Concurrence, for implementing a production plan without government support (Conseil de la Concurrence, 1998). The plan consisted of imposing penalties on individual producers for production in excess of a predetermined quota. A similar plan was approved by the government immediately afterwards, and such production plans are still in place today. The authority's 1998 ruling stated:[2]

> [...] it remains undisputed that, on the one hand, the [supply control] measure targeted all Comté-producing firms and, on the other hand, [...] the criticised measure had a deterrent effect that limited the price decrease and made Comté production less attractive for Emmental producers wishing to shift to Comté production; [...] finally, that several firms were indeed charged a penalty for producing above their allocated quota, for a total amount of 1,156,509 French Francs; that, as a result, the measure had a significant effect on the market in question [...].

In a 2000 report, the Organisation for Economic Cooperation and Development expressed concerns regarding the existence of market power associated with certain European PDOs, explicitly referring to specialty cheese markets (Lucatelli, 2000). In addition, approval of the Comté production plan by public authorities, in particular, the French Ministry of Finance, has sometimes been difficult to obtain.[3]

Nonetheless, the empirical importance of the existing distortion remains unknown. An empirical analysis that estimates the extent of market power being exercised in the Comté market is thus of interest to policymakers and potentially to antitrust authorities. More generally, economists have been prompt in describing European PDOs as cartelised markets. This argument may be used to oppose the recognition of PDOs at the international level.[4] Although focused on a particular commodity, this study sheds doubt on the ability of collective marketing arrangements such as those observed in certain PDO markets to sustain supercompetitive prices.

Traditionally, the NEIO methodology has been used to measure the effects of industry concentration on buyer or seller competition and the efficiency of markets.[5] It has also been used, in agricultural applications, to test whether marketing institutions such as producer cooperatives or marketing orders benefiting from antitrust exemptions have been successful in extracting oligopoly rents (Crespi and Sexton, 2003; Crespi and Chacón-Cascante, 2004). The present paper belongs to this latter branch of literature.

2 Translated from French by the author.

3 A Comté-producing firm brought a case to the French antitrust authority in late 2006 regarding potential antitrust infringements contained in a proposed amendment to the Comté regulations (Conseil de la Concurrence, 2007). The new provisions, which have since been adopted by the French government, include several limitations on the production capacity of Comté-producing entities.

4 For an overview of the debate over geographical indications between the EU and the US, see, for instance, Josling (2006).

5 For a review of NEIO studies in the food sector, see Kaiser and Suzuki (2006).

This article addresses the concern that the NEIO methodology has often been implemented on overly aggregated industries (Kaiser and Suzuki, 2006). By focusing on one cheese variety, without ignoring the possibility of substitution with other cheeses, the study narrowly defines the potentially imperfectly competitive industry and tailors the estimation procedure to a close observation of policy, technology and demand conditions.

The existence of a pre-existing distortion due to the EU dairy quota policy is taken into consideration by specifying cost as the opportunity cost of not producing a substitute, namely Emmental cheese.[6] This constitutes an interesting adaptation of the standard NEIO model. A fundamental reason for defining the cost of producing Comté as the opportunity cost of not producing Emmental is that this opportunity cost includes any dairy quota rent accruing to the dairy production sector and is therefore a better indication of the true cost of producing Comté than a sum of physical input costs. In fact, ignoring the possibility that the dairy quota has been binding over the period of investigation could lead the analyst to spuriously attribute a significant price–cost margin to the presence of market power at the level of the Comté industry, while such a margin could be due, in part or in total, to the dairy quota, and exist in other dairy markets as well. By specifying marginal cost as the opportunity cost of not producing Emmental, we thus seek to detect any *additional* mark-up beyond that originating in the European dairy quota.[7]

The estimation is performed on quarterly data at the wholesale stage over the period 1985–2005. It leads to the conclusion that if market power has been exerted by the Comté industry, it is hard to detect econometrically and is likely to be very small. This finding suggests that consumers have not been hurt by supply control and that the social cost of the policy has been negligible.

2. The French Comté cheese market

Comté is a pressed, cooked cheese made from raw cows' milk, aged for at least four months, which comes in large wheels weighing between 66 and 106 lb. The specificity of Comté cheese was recognised by a Court decision in 1952, and its production was first codified in 1958. Comté was introduced in the European register of protected designations of origin in 1996, the date the register was created (Council of the European Union, 2006). Production of Comté is currently regulated by a 2007 governmental decree (République

6 Emmental cheese is the main type of pressed, cooked cheese manufactured in France, with an annual production of about 250,000 metric tons.

7 This study is not the first one to suggest using the price of a close substitute in the marginal cost specification. Arnade and Pick (1999) pointed out that when a single commodity is sold in two different markets, one can use the *difference* in price in the two markets to identify the degree of competitiveness of each market. Provided that the cost of production is the same for the two markets, marginal cost terms cancel out [see equation (3) of their article]. Our approach differs from theirs in that (i) Comté and Emmental have different production costs, so that we need to keep a 'cost difference' term in the opportunity cost specification and (ii) the market for Emmental is assumed to be competitive, so that we estimate only one conduct parameter.

Française, 2007). The decree contains provisions such as the delimitation of the area of production, the physical characteristics of the cheese and restrictions on the production methods to be used at each stage of fabrication, including the farm level. Notable production constraints include restrictions on cow breeds and feed, and limits on the stocking rate and the distance travelled for collecting milk.

The production process unfolds in three stages: milk production, cheese fabrication and cheese ripening. All stages must take place within a delimited geographical area covering several districts of the Franche-Comté region.[8] In January 2005, there were about 3,300 milk producers, 190 cheese factories and 20 ripening facilities involved in Comté cheese production. Notably, 85 per cent of the cheese factories were owned by milk producers through cooperatives (Comité Technique du Comté, 2005).

Milk producers, dairy cooperatives and ripening facilities are represented by a producer association, CIGC, whose stated missions are to guarantee the specificity of Comté cheese and help producers maintain a sustainable activity in the region. CIGC fulfils its first mission by controlling producers at various stages of the production process, filing lawsuits against imitators and participating in the development of standards.

The second mission is fulfilled through technical assistance to producers, generic advertising and the promotion of Comté cheese in export markets. CIGC also makes a yearly production plan to limit the quantity of Comté cheese produced, paired with model contracts designed to redistribute income between milk producers and ripening facilities.

The production plan is enforced through the delivery of certification marks necessary to authenticate cheese wheels. Marks are purchased from CIGC by cheese factories and applied onto unripe cheese wheels. Each factory is allocated a quota and charged a penalty for each mark purchased beyond that quota. The penalty is adjusted to account for any modification in the average weight of wheels.[9] The production plan is subject to yearly approval by the government.[10]

The main purpose of the model contracts, to be adopted by ripening facilities and the cheese factories supplying them, is to set the price of unripe cheese.[11] The contract also compels the ripening facility to purchase all the cheese produced by its suppliers.

Therefore, the Comté production sector can be modelled as a vertically integrated entity, with CIGC choosing the total quantity to be produced and reallocating rents through the contractual price of unripe cheese. These rents

8 The geographical constraint is not binding. Between 60 and 70 per cent of the milk produced in the eligible region is transformed into Comté cheese.

9 Colinet *et al.* (2006) estimated the over-quota premium to be about 36 per cent of the wholesale price of Comté cheese in 2006.

10 The production plan is approved by the *Office National du Lait et des Produits Laitiers*, the Ministry of Agriculture and the Ministry of Finance.

11 According to the Director of CIGC, about 40 per cent of ripening facilities have formally signed the model contract. However, the vast majority of ripening facilities follow the price specified in the contract, even if they have not signed it.

include the quota rent from the EU dairy quota scheme, as well as any additional rent generated by the supply control scheme at the expense of buyers of ripened cheese. In what follows, we refer to the vertically integrated Comté production sector as 'the Comté industry'.

3. The theoretical model

Assume that producers in the Comté region can produce two goods, Comté cheese and Emmental cheese,[12] according to the aggregate cost function $C(Q_{\mathrm{C}}, Q_{\mathrm{E},-2})$, where Q_{C} and $Q_{\mathrm{E},-2}$ denote the quantities of Comté and Emmental produced. The two-period lag on the Emmental quantity reflects the fact that Comté cheese is ripened for a longer period than Emmental cheese. We assume that Emmental is sold in the same quarter it is produced, whereas Comté is ripened for two extra quarters.[13] Therefore, the function C represents the total cost, as of the current period, of producing Q_{C} units of Comté to be sold in the current period and $Q_{\mathrm{E},-2}$ units of Emmental to be sold two periods earlier. It includes all costs, from the production of milk to the ripening of cheese. The function C is assumed to have all desirable properties, in particular, it is non-decreasing in its arguments and convex. Producers are endowed with $\bar{M}$ milk quotas, and the milk quota is binding.[14] Although producers located in the Comté region are the sole producers of Comté cheese, they produce only a small share of the total milk supply, and therefore we assume that they cannot influence the price of generic milk.[15] Entry into the Emmental market is free.[16] We assume that the transformation of milk into Emmental cheese involves a fixed-proportion, constant-returns-to-scale

12 The share of Comté and Emmental among the four main cheeses produced in Franche-Comté (Comté, Emmental, Morbier and Raclette) was 96.5 per cent at the beginning of the study period and decreased to 83.9 per cent in 2005. Over the period, the mean was about 89.8 per cent (Ministère de l'Agriculture et de la Pêche, 2007b). These shares are rough estimates computed by comparing the quantities of each cheese variety, and do not account for differences in their milk content. It is expected that the milk content of Comté and Emmental is higher than that of the other two cheeses, which are not cooked. Notably, the vast majority of new entrants in the Comté industry over the period of investigation have originated in cheese factories previously specialised in Emmental production (Comité Interprofessionnel du Gruyère de Comté, 2006; Ministère de l'Agriculture et de la Pêche, 2007a).

13 Over the period, Comté cheese was ripened about 5.5 months longer than Emmental, that is, about two quarterly periods. Emmental cheese is ripened for a minimum of 6 weeks.

14 EU dairy quota appears to have been binding in the Comté region. While milk deliveries in Franche-Comté had been increasing at a rate of over 1 per cent per year in the decade preceding the quota policy, this trend stopped with the introduction of quota in 1984. Milk quota for 2004–2005 was 1 per cent below the average of milk deliveries over the period 1981–1983 in this region (Ministère de l'Agriculture et de la Pêche, 2005). Colinet *et al.* (2006) reported that between 1992 and 2004, milk quota in the Comté eligible area (Departments of Doubs and Jura) has also limited milk deliveries.

15 Milk deliveries in the Franche-Comté region represented about 4.8 per cent of all French milk deliveries in 2005 (Ministère de l'Agriculture et de la Pêche, 2007b).

16 Emmental cheese can be produced anywhere. Today, more than 70 per cent of the total production of Emmental in France occurs outside the traditional area of production. France both imports and exports Emmental cheese. In 2004, France exported 53,600 tons of Emmental and imported 36,466 tons, mostly from EU countries (Syndicat Interprofessionnel du Gruyère Français, 2004).

technology, so that the price of Emmental, which is then equal to the price of generic milk plus a unit processing cost, does not depend on the quantities of Comté (Q_C) or Emmental ($Q_{E,-2}$) produced in the Comté region.

Denote by k_C the coefficient of conversion of milk into Comté cheese and by k_E the coefficient of conversion of milk into Emmental cheese, both assumed to be fixed. That is, one unit of dairy quota can be used to produce k_C units of Comté cheese or k_E units of Emmental cheese. The inverse demand function for Comté cheese is denoted by $P(\cdot)$, and $P_{E,-2}$ denotes the price of Emmental in period -2. To capture the opportunity cost of time, we assume that the revenue from selling Emmental two periods earlier than Comté can grow at the interest rate INT_{-2}, expressed in per cent at 6 months. We denote by $\tilde{P}_{E,-2}$ the lagged price of Emmental as of the current period, that is, $\tilde{P}_{E,-2} = P_{E,-2}\,(1 + \text{INT}_{-2}/100)$.

CIGC is assumed to behave as a tentative monopolist seeking to maximise net returns to milk quota ownership in the Comté region.

We first assume that CIGC acts as a joint profit-maximising cartel, subject to the quota constraint. It is further assumed that CIGC has perfect foresight of the future demand conditions in the Comté market and knows the cost function C. CIGC's allocation problem in period -2 can thus be written as

$$\max_{Q_C} \quad P(Q_C)Q_C + \tilde{P}_{E,-2}k_E\left(\bar{M} - \frac{Q_C}{k_C}\right) - C\left(Q_C, k_E\left(\bar{M} - \frac{Q_C}{k_C}\right)\right). \tag{1}$$

The first-order condition is

$$P(Q_C) + P'(Q_C)Q_C - \frac{k_E}{k_C}\tilde{P}_{E,-2} - C_1 + \frac{k_E}{k_C}C_2 = 0, \tag{2}$$

where C_i, $i \in \{1,2\}$ denotes the first derivative of C with respect to its ith argument. This optimisation condition can be rewritten as

$$\text{MR}(Q_C) = \frac{k_E}{k_C}\tilde{P}_{E,-2} + \phi(Q_C), \tag{3}$$

where MR denotes the marginal revenue function for Comté and

$$\phi(Q_C) = C_1\left(Q_C, k_E\left(\bar{M} - \frac{Q_C}{k_C}\right)\right) - \frac{k_E}{k_C}C_2\left(Q_C, k_E\left(\bar{M} - \frac{Q_C}{k_C}\right)\right).$$

It is easy to show from the convexity of C that ϕ is a non-decreasing function of Q_C.[17]

Suppose now that CIGC does not exercise its full monopoly power, so that, following the NEIO paradigm, the market-level, imperfectly competitive

17 See the Appendix for the derivation.

equilibrium is described by an equation of the type

$$\mathrm{PMR}(Q_C) = \frac{k_E}{k_C}\tilde{P}_{E,-2} + \phi(Q_C), \tag{4}$$

with $\mathrm{PMR}(Q_C)$ denoting the *perceived* marginal revenue schedule. Following Bresnahan (1982), it is convenient to write the function PMR (Q_C) as a convex combination of the market inverse demand $P(Q_C)$ and the true marginal revenue $\mathrm{MR}(Q_C)$, that is,

$$\mathrm{PMR}(Q_C) = (1-\theta)P(Q_C) + \theta\,\mathrm{MR}(Q_C) \tag{5}$$

for some $\theta \in [0,1]$. The weighting parameter θ can then be interpreted as an overall degree of non-competitiveness, as suggested in Hyde and Perloff (1995).[18] It is referred to in the NEIO literature as the market power or conduct parameter, and is equal to zero if the industry is competitive and to 1 if the industry behaves as a monopolist. Intermediate values of θ correspond to market equilibria between perfect competition and pure monopoly. The parameter θ is directly related to the Lernex index L of imperfect competition for this industry, $\theta = L|\eta_{CC}|$, where η_{CC} denotes the own-price elasticity of demand for Comté cheese.

Equations (4) and (5) can be combined and rearranged to yield the pricing equation

$$P_C\left(1+\frac{\theta}{\eta_{CC}}\right) = \frac{k_E}{k_C}\tilde{P}_{E,-2} + \phi(Q_C), \tag{6}$$

where P_C denotes the price of Comté. The coefficient k_E/k_C on the lagged price of Emmental reflects the difference in milk content between the two cheese varieties, whereas the non-decreasing function $\phi\,(Q_C)$ represents the marginal cost difference between Comté and Emmental (i.e. the cost of producing one more unit of Comté and k_E/k_C fewer units of Emmental).

The idea behind the cost specification on the right-hand side of equation (6) is that, by producing one unit of Comté cheese, the industry forfeits the rent that it would earn if it produced and sold the quantity of Emmental cheese corresponding to the amount of EU milk quota used to produce this unit. Therefore, in equilibrium, the industry equates the perceived marginal revenue from producing an additional unit of Comté cheese to the marginal opportunity cost of doing so, which is equal to the marginal revenue forfeited from potential sales of Emmental cheese plus the difference in marginal costs, $C_1 - (k_E/k_C)C_2$. This opportunity cost specification assumes that the milk content of both cheese varieties is fixed and does not depend on the quantities produced, an assumption parallel to the traditional fixed-proportions assumption ubiquitous in the NEIO literature. Given that slightly more milk is used to produce 1

18 This interpretation is also discussed by Kaiser and Suzuki (2006: 21).

kg of Emmental than 1 kg of Comté, the coefficient k_E/k_C should be close to, but less than, 1. More precisely, the productivity of milk in Comté is about 10 per cent (i.e. 10 kg of Comté can be made from 100 kg of milk), whereas that of Emmental ranges between 8 and 9 per cent. Thus, k_E/k_C is expected to be between 0.8 and 0.9.[19]

4. Estimation strategy

Using time-series data on the price and quantity of Comté cheese, we estimated equation (6), together with the demand equation, in order to determine jointly the demand parameters, the parameters of the opportunity cost function and the market power parameter, θ. As is apparent from equation (6), identification of θ relies on temporal variation in the demand elasticity, and so we need to introduce demand shifters that also shift the demand elasticity.

4.1. Demand

The elasticity shifters included in the demand equation are the price of Emmental cheese, income and quarterly dummies. By 'elasticity shifters', we mean variables that interact with the price of Comté cheese on the right-hand side of the demand equation (quantity being the dependent variable), and therefore allow the calculated demand elasticity to vary across time periods.[20] To our knowledge, this is the first NEIO study to incorporate three or more different elasticity shifters in the demand equation.[21]

The demand for Comté cheese is specified as

$$q_C = \alpha_1 + \sum_{i=2}^{4} \alpha_i S_i + \left(\beta_1 + \sum_{i=2}^{4} \beta_i S_i + \beta_E P_E + \beta_I I \right) P_C + \gamma P_E + \delta I + e_d, \quad (7)$$

where q_C denotes the per capita quantity of Comté cheese sold, S_i are quarterly dummy variables for the 2nd, 3rd and 4th quarters, respectively, P_C and P_E are the real wholesale prices of Comté and Emmental, respectively, and I denotes the real per capita net disposable income. Prices and income are adjusted for inflation using the general consumer price index.

This demand specification has the desirable property of allowing each of the shifters (price of Emmental, income and quarter) to change demand additively *and* change the demand slope. This flexibility is necessary to avoid spurious effects of those shifters on the implied demand elasticity.

19 Colinet *et al.* (2006) assumed that 10.8 and 12 litres of milk are required to produce 1 kg of ripened Comté and Emmental, respectively. This would correspond to a value of k_E/k_C equal to 0.9 exactly.

20 In addition, the demand elasticity may vary due to functional form assumptions.

21 Typically, shifters are included either as interaction terms with price ('slope shifters') or as additive terms ('intercept shifters'). This practice unduly restricts the way demand depends on each shifter and could lead to spurious effects of shifters on the calculated elasticity. Crespi and Chacón-Cascante (2004) discussed the choice of demand shifters in existing NEIO studies.

To assess the importance of functional form assumptions, we also estimated a log–log version of equation (7):

$$\log q_{\mathrm{C}} = \alpha_1 + \sum_{i=2}^{4} \alpha_i S_i + \left(\beta_1 + \sum_{i=2}^{4} \beta_i S_i + \beta_{\mathrm{E}} \log P_{\mathrm{E}} + \beta_I \log I\right) \log P_{\mathrm{C}}$$
$$+ \gamma \log P_{\mathrm{E}} + \delta \log I + u_{\mathrm{d}}. \tag{8}$$

4.2. Supply

The supply relationship is given by equation (6). We further assume that the function $\phi(Q_{\mathrm{C}})$ is linear in Q_{C}, so that the equation used for estimation is

$$P_{\mathrm{C}}\left(1 + \frac{\theta}{\eta_{\mathrm{CC}}}\right) = c_0 + c_1 P_{\mathrm{E},-2}\left(1 + \frac{\mathrm{INT}_{-2}}{100}\right) + c_2 Q_{\mathrm{C}} + e_{\mathrm{s}}, \tag{9}$$

where the demand elasticity η_{CC} is a function of the demand parameters β_i ($i \in \{1, 2, 3, 4\}$), β_{E} and β_I derived from specification (7) or (8) and c_0, c_1 and c_2 are cost parameters to be estimated.[22]

The prices P_{C} and $P_{\mathrm{E},-2}$ are expressed in real terms adjusted for inflation and therefore the interest rate INT_{-2} represents the real interest rate, also adjusted for inflation. This real interest rate is computed from the equation $1 + \iota = (1 + \mathrm{INT})(1 + \pi)$, where ι denotes the nominal interest rate and π the rate of inflation (Perloff, 2009).

The presence of the term $c_2 Q_{\mathrm{C}}$ on the right-hand side of equation (9) allows us to test for scale in the conversion from Emmental production to Comté production. In particular, if farms were heterogeneous in their ability to shift from Emmental to Comté production, we would expect the coefficient c_2 to be positive. As argued in Section 3, the coefficient c_1 should reflect the difference in milk content between Comté and Emmental, and is expected to range between 0.8 and 0.9.

4.3. Choice of instruments

The contemporaneous and lagged prices of Emmental cheese were used as instruments in the empirical estimation procedure. The assumption that these prices are exogenous to the demand equation (7) or (8) and the supply relation (9) is justified by the fact that entry is free in the Emmental market, so that the Emmental market can be assumed to be competitive. In addition, if generic milk is transformed into Emmental cheese according to

22 Lagged values of the interest rate were added as explanatory variables in the empirical specification of $\phi(Q_{\mathrm{C}})$, to capture the fact that some costs are incurred in earlier periods. The associated coefficients had the wrong sign and were not statistically significant; therefore, it was decided to drop those variables from the cost specification.

a fixed-proportion, constant-returns-to-scale technology, as assumed, then the exogeneity of the Emmental price is satisfied as long as random shocks to the demand for Comté cheese and to its supply relation are uncorrelated with the equilibrium price of generic milk, which under the milk quota system is determined by the available quota and the total milk demand. These random shocks in the Comté market may potentially affect the equilibrium price of generic milk in two ways: by shifting the total supply of generic milk or the demand for dairy products other than Comté cheese. We argue that both effects should be negligible, given that the milk market can be considered European-wide. First, even though random shocks to the demand for Comté or its supply relation affect the quantity of milk quotas used for Comté production, and therefore that available to produce generic milk, this is unlikely to affect the total milk supply, since Comté represents about 2 per cent of the total milk collected in France. Second, random shocks to the price of Comté may shift the demand for substitutes of Comté, but again this is unlikely to affect the derived demand for milk (and therefore its equilibrium price), because substitutes for Comté cheese represent a relatively small share of total milk use in France and in the EU.

Other instruments include the seasonal dummies and the income variable.

4.4. Data frequency

Most of the variables needed, except income and population, are available with a monthly frequency. Since the production plan sets production caps for a one-year period, it seems to call for the use of annual data. Such an approach, however, would not only considerably reduce the sample size (in the present case, to 21 observations), but would also ignore specific provisions of the production plan. First, the plan can be revised during the year (République Française, 1977). Second, CIGC can adopt exceptional compulsory or voluntary measures, such as withdrawals of eligible milk or unripe cheese wheels, whenever the market situation is deemed unfavourable. Although compulsory withdrawals have been exceptional, voluntary (but financially encouraged) withdrawals have been used more frequently. The possibility of adjustments to the production plan during the year may therefore partly justify the use of higher frequency data.

Another consideration when deciding upon data frequency is the observed differences in the ripening time of Comté cheese. Ageing varies from 4 to 24 or even 30 months. In addition, cheese wheels can be stored at low temperatures to suspend the ripening process, which further increases firms' ability to delay the marketing of cheeses. Therefore, one cannot rule out the endogeneity of the ripening time. The strategic decisions of ripening facilities with respect to ripening and storage are ignored by using a static model of imperfect competition and assuming that the industry is vertically integrated. The error associated with using a static framework is likely to be greater if high-frequency data are used, whereas it would theoretically disappear if the frequency was low enough to assume that all cheese wheels are ripened within one period.

Taking all these considerations into account, we chose to use quarterly data. Although this does not totally solve the issue of endogenous ripening, it should be less critical than with monthly data, especially because the bulk of the production is sold between 6 and 8 months of age, with a tendency towards longer ripening times at the end of the period of investigation.[23]

5. Data description

The data cover the period from the first quarter of 1985 to the fourth quarter of 2005. The starting date was chosen as 1985 because the EU milk quota scheme was introduced in April 1984. Industry data on the wholesale price and marketed quantity of Comté cheese comes from CIGC. It is not possible to distinguish between domestic and export sales. However, given the very small share of exports (about 5 per cent), total quantity should be a reasonable proxy for sales in the French market. Industry data on the wholesale price of Emmental cheese comes from SIGF (*Syndicat Interprofessionnel du Gruyère Français*).[24] The net disposable national income for France was obtained from Eurostat, as was the population variable used to construct the per capita Comté cheese consumption and net disposable income variables. The general CPI for France was obtained from the OECD database, and the nominal interest rate from the *Banque de France* webpage. Summary statistics are reported in Table 1.

6. Estimation and results

The simultaneous equation system consisting of the demand equation (7) or (8) and the pricing equation (9) was estimated using the iterative non-linear optimal generalised method of moments (Hansen, 1982).

Instruments for equation (7) or (8), denoted by the row vector Z_{d}, include a constant, the logarithms of P_{E} and I, seasonal dummies, and a series of interaction terms constructed from $\tilde{P}_{\mathrm{E},-2}$ to instrument for each of the endogenous regressors involving the price of Comté. Instruments for equation (9), denoted by the row vector Z_{s}, include a constant, $\tilde{P}_{\mathrm{E},-2}$, the logarithms of P_{E} and I, and seasonal dummies. The weighting matrix is constructed allowing for heteroscedasticity and autocorrelation up to three lags and assuming that the error terms e_{d} (u_{d}) and e_{s} are uncorrelated.

23 To test the sensitivity of estimates to the choice of data frequency, a model using monthly data was also estimated. Income, which was available only at a quarterly frequency, was distributed over the months within each quarter. Estimates were close to those obtained using quarterly data, and conclusions regarding market power were unaffected.

24 For both Comté and Emmental, some adjustments to the raw data were necessary to account for changes in the way the industry price was calculated over the period of investigation. Detailed information regarding these adjustments is available in the Appendix.

Table 1. Summary statistics: sample means (standard deviations in parentheses)

Variable (units of measurement)	Full sample	First quarter	Second quarter	Third quarter	Fourth quarter
Total quantity (Q_C) (1,000 tons)	9.08 (1.49)	8.75 (1.36)	8.56 (1.36)	9.08 (1.37)	9.93 (1.59)
Per capita quantity (q_C) (100 g per head)	1.52 (0.212)	1.47 (0.188)	1.43 (0.187)	1.52 (0.187)	1.66 (0.221)
Real price of Comté (P_C) (€/kg, 2000 monetary values)	5.58 (0.378)	5.59 (0.417)	5.59 (0.403)	5.56 (0.360)	5.57 (0.356)
Real price of Emmental (P_E) (€/kg, 2000 monetary values)	4.99 (0.367)	5.03 (0.409)	4.99 (0.364)	4.96 (0.340)	4.96 (0.375)
Real income (I) (thousand euros, 2000 monetary values)	4.60 (0.563)	4.57 (0.574)	4.58 (0.575)	4.61 (0.576)	4.64 (0.565)
Real interest rate (INT_{-2}) (per cent, 6-monthly rate)	3.08 (1.29)	–	–	–	–
Number of observations	84	21	21	21	21

For the system consisting of (7) and (9), we denote

$$e = \begin{pmatrix} e_\mathrm{d} \\ e_\mathrm{s} \end{pmatrix} \quad \text{and} \quad Z = \begin{pmatrix} Z_\mathrm{d} & 0 \\ 0 & Z_\mathrm{s} \end{pmatrix}.$$

The moment conditions used for estimation are

$$E(e|Z) = 0, \tag{10}$$

where E denotes the expectation operator. The vector of unknown model parameters is denoted by $\lambda = (\alpha_1, \alpha_2, \alpha_3, \alpha_4, \beta_1, \beta_2, \beta_3, \beta_4, \beta_\mathrm{E}, \beta_\mathrm{I}, \gamma, \delta, c_0, c_1, c_2, \theta)$. Using bold letters to denote the sample equivalents of each random

Table 2. Parameter estimates obtained from the iterated NLOGMM estimation

Parameters	Linear model	Log–log model
α_1	−13.198* (3.776)	−26.862* (7.893)
α_2	0.082 (0.129)	0.184 (0.131)
α_3	0.386* (0.170)	0.405 (0.211)
α_4	0.713*(0.235)	0.408 (0.294)
β_1	2.446* (0.676)	17.991* (4.582)
β_2	−0.021 (0.023)	−0.121 (0.076)
β_3	−0.061* (0.029)	−0.218 (0.121)
β_4	−0.094* (0.042)	−0.166 (0.171)
β_E	−0.162* (0.053)	−4.596* (1.252)
β_I	−0.418* (0.110)	−7.730* (1.933)
γ	0.937* (0.312)	7.984* (2.214)
δ	2.506* (0.592)	13.783* (3.260)
c_0	0.959 (0.643)	1.004 (0.671)
c_1	0.821* (0.092)	0.808* (0.103)
c_2	0.037 (0.034)	0.039 (0.033)
θ	0.001 (0.010)	0.002 (0.007)
Comté (F(6,152))	3.360*	2.827*
Second quarter (F(2,152))	5.593*	6.059*
Third quarter (F(2,152))	7.706*	7.115*
Fourth quarter (F(2,152))	57.833*	54.334*
Emmental (F(2,152))	4.737*	7.219*
Income (F(2,152))	84.481*	69.041*
J-test ($\chi^2(3)$)	0.333	0.344
$\bar{\eta}_\mathrm{CC}$	−1.202* (0.294)	−1.249* (0.313)
$\bar{\eta}_\mathrm{CE}$	0.109 (0.127)	0.093 (0.145)
$\bar{\eta}_\mathrm{CI}$	0.531* (0.091)	0.511* (0.091)

Standard errors are indicated in parentheses. The reported F-statistics relate to the joint significance of regressors that include the variable of interest. The third part of the table reports statistics for the J-test of overidentifying restrictions. The last part reports the implied own-price, cross-price and income elasticities of demand at the mean of the sample.

*Statistically significant at the 5 per cent level.

variable, the estimates $\hat{\lambda}$ were chosen so as to minimise

$$Q_{\mathrm{T}} = \frac{1}{T}\mathbf{e}'\mathbf{Z}\hat{\mathbf{S}}^{-1}\mathbf{Z}'\mathbf{e},$$

where T denotes the sample size. The weighting matrix $\hat{\mathbf{S}}$ was estimated recursively as the HAC Newey–West estimate of

$$S = \frac{1}{T}\sum_{i=1}^{T}\sum_{i=1}^{T} E\left[Z_i' e_i e_j' Z_j\right].$$

The results from estimating parameters λ using the linear and double-log demand specifications are displayed in Table 2. Reported standard errors are heteroscedasticity-robust and corrected for autocorrelation with three lags according to the Newey and West (1987) procedure. A three-lag specification is chosen because the production plan is adopted on a yearly basis, and therefore the quantity marketed in each quarter is likely to be correlated with that marketed in the other quarters of the same year. Hansen's (1982) J-test of over-identifying restrictions fails to reject the model or the set of stochastic assumptions (10) used for identification. In particular, the exogeneity of the contemporaneous and lagged prices of Emmental to demand and the supply relation cannot be empirically rejected.

6.1. Demand estimates

The implied own-price, cross-price and income elasticities, evaluated at the mean of the sample, all have the expected signs and reasonable magnitudes, and are consistent between the two models.[25] The own-price and income elasticities of demand are statistically significant in the two models. Although the implied cross-price elasticity is not significant at the sample mean, we can reject the null hypothesis that $\beta_{\mathrm{E}} = \gamma = 0$ in the two models, at the 5 per cent significance level.

The implied own-price elasticity of demand at the sample mean is -1.202 for the linear model and -1.249 for the double-log model, which seem reasonable given available elasticity estimates for the entire cheese group in France (e.g. -0.83 according to Combris *et al.* (1998); -0.648 for the compensated demand elasticity (Nichèle, 2003)). These figures are also consistent with the estimate of -1.28 reported in Colinet *et al.* (2006) for the own-price elasticity of demand for Comté cheese. The temporal variation in the own-price elasticity of demand is shown in Figure 1, and is consistent between the two models. The secular trend in the absolute elasticity of demand has been increasing over the 20-year period, with a slight downward adjustment

25 The magnitude of the income elasticity of demand is not consistent with the general belief that Comté cheese is a luxury good. This may be due to the parsimony of the demand specification and the use of total income rather than expenditure on a more narrowly defined group.

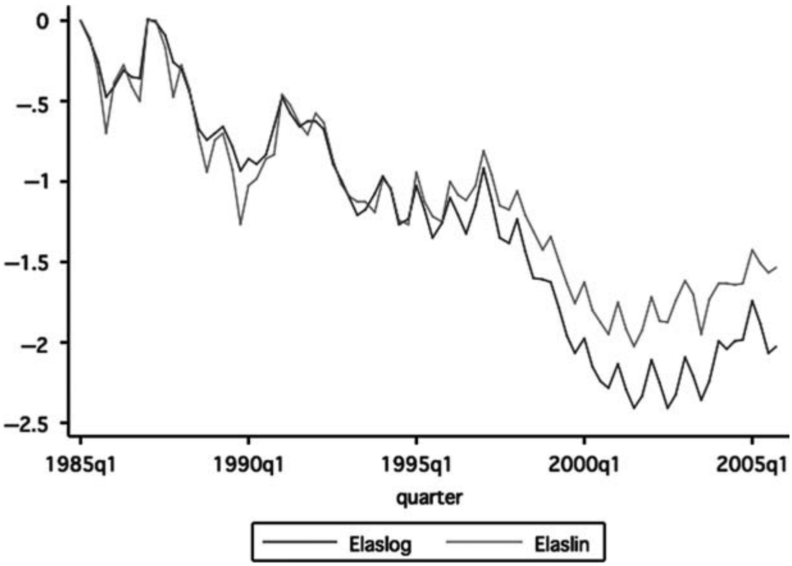

Figure 1. Own-price elasticities derived from the log–log (elaslog) and linear (elaslin) models.

in the last 3 years. Overall, demand for Comté appears to have become more elastic over time.

All demand shifters included in the demand specifications (7) and (8) are significant, based on the *F*-tests reported in the second part of Table 2. Therefore, seasonality, the price of Emmental and the income variable all seem to affect demand. Although the separate effects of the Emmental price and income shifters on the intercept and the slope of demand are statistically significant, this is not always the case for the quarterly shifters taken individually. However, when tested for their joint effect on the intercept and the slope, all three seasonal effects taken individually are highly significant, particularly the fourth-quarter effect. Furthermore, the signs of the seasonal effects are consistent across the two tested models. In particular, demand intercepts are larger and demand slopes steeper, during the last three quarters, compared with the baseline first quarter. More interestingly, we confirm statistically that demand for Comté cheese reaches a peak during the fourth quarter, a trend that is well understood by industry representatives and can be attributed to increased consumption during the Christmas holiday period. This conclusion can be drawn by conducting pairwise tests that compare, for any two seasons, the *combined* effects of the slope and intercept seasonal dummies on the quantity demanded, the price of Comté being set at its sample mean.[26]

26 For instance, to see whether demand was significantly higher in the fourth quarter than in the third, other factors being held constant, we tested the hypothesis

$$H_0\colon \alpha_4 + \beta_4\bar{P}_C - (\alpha_3 + \beta_3\bar{P}_C) = 0 \text{ against } H_1\colon \alpha_4 + \beta_4\bar{P}_C - (\alpha_3 + \beta_3\bar{P}_C) > 0.$$

6.2. Cost and conduct estimates

Cost and conduct estimates seem plausible and are only marginally sensitive to the demand specification.

The coefficient of the lagged price of Emmental cheese, c_1, is highly significant and is smaller than 1 under the two demand specifications. We can reject the hypothesis that $c_1 = 1$, based on the test $c_1 = 1$ vs. $c_1 < 1$, at the 5 per cent significance level. The point estimates are within the expected range of [0.8, 0.9] and the value of 0.9 lies within the 95 per cent confidence interval.

The p-values for the estimates of coefficient c_2 (0.278 for the linear specification and 0.240 for the double-log specification) imply that the hypothesis of constant returns to scale in the conversion from Emmental to Comté cannot be rejected, at the 5 per cent significance level. In terms of the function ϕ, this means that we fail to reject the hypothesis that $d\phi/dQ_C = 0$, at least locally. Given that the milk processing and ripening stages for the two cheeses are more likely to involve constant returns to scale than farm operations, this result may be an indication of a certain farm homogeneity among those producers who have shifted from Emmental to Comté over the period of investigation. In other words, the result could mean that farm heterogeneity is not sufficient, around the observed equilibrium, to link the observed increase in the quantity of Comté cheese produced to higher production costs.

Finally, and most importantly, given the purpose of this article, the market power estimate is positive, small and statistically insignificant in both models. The null hypothesis of perfect competition in the Comté market cannot be rejected at the 5 per cent significance level, based on the test $\theta = 0$ vs. $\theta > 0$. In contrast, monopoly is easily rejected, based on the test $\theta = 1$ vs. $\theta < 1$. Market power estimates are well within values traditionally considered to be close enough to perfect competition not to raise economic efficiency concerns (Sexton and Zhang, 2001).[27]

7. Conclusion

In this article, an NEIO approach was used to measure the degree of seller market power in the French Comté cheese industry, characterised by vertical contracts between milk producers and ripening facilities, and a government-sanctioned supply control scheme that has caught the attention of the French antitrust authority. The inclusion of a set of elasticity shifters in the

27 We also estimated a modified model that allowed the market power intensity to take different values for the two sub-periods 1985q1–1995q3 and 1995q4–2005q4. The choice of sub-periods was motivated by the observation that production caps up to 1994–5 had little effect in practice. First, Comté production was not attractive in the late 1980s, so that entry into the industry did not have to be prevented. Second, even though the attractiveness of the Comté sector increased in the early 1990s, the over-quota penalty was too small to discourage entry. The penalty was increased in 1995–6 and remained at high levels thereafter. Given the ripening time of Comté cheese, if binding caps took hold in March 1995, the effect should be detectable starting two periods later. Although the point estimate of the conduct parameter was higher for the second sub-period than for the first, both estimates were statistically insignificant. In addition, the hypothesis that the two conduct parameters are equal could not be rejected (p-value$= 0.808$).

demand specification permitted identification of the market power parameter. One original feature of the model lies in the cost specification, which includes the price of Emmental cheese, to take account of the EU milk quota policy and the substitutability between the two cheeses in production. Since milk quota is likely to create a wedge between the marginal cost of milk and its market price, defining the opportunity cost of producing Comté in terms of the net revenue forgone from potential sales of Emmental, rather than a sum of input expenses, is critical to avoid attributing any significant price–cost margin solely to market power exercised at the level of the Comté industry.

The hypothesis of perfect competition could not be rejected. In contrast, monopoly was clearly rejected. Elasticity, cost and market power estimates were robust to the demand specification.

This study sheds serious doubt on the ability of the observed supply control scheme to allow the Comté cheese industry to exert significant market power towards buyers. Small values of the market power parameter imply that the associated deadweight loss has been negligible over the period. This conclusion contrasts with the assumptions underlying the ruling of the French antitrust authority and provides some reassurance regarding the efficiency of PDO markets that allow collective marketing arrangements.

A possible interpretation of the lack of measurable market power in the Comté market is that the control exercised by the governmental entities in charge of approving the yearly production plan effectively prevents CIGC from restricting supply far from the competitive level. Unless explicitly approved through an administrative procedure that involves the ministerial departments in charge of agriculture and competition, the production plan is not legally enforceable by CIGC and becomes subject to scrutiny by the *Conseil de la concurrence* (République Française, 1977). While it is difficult to infer, from purely anecdotal evidence, the extent of governmental control exercised through this administrative process over the period of investigation, it is noteworthy that the *Conseil* wrote, in its 1998 ruling:[28]

> In the period preceding the decision of 4 March 1995, [CIGC], having acknowledged in a decision of 6 July 1993 the refusal of the Ministry of Finance to approve a production limitation,[29] had therefore deemed necessary to stimulate sales through promotion.

The fact that the Ministry of Finance explicitly refused any production limitation in 1993 suggests that later approvals were possible only because the proposed quantity was sufficiently close to the competitive quantity.[30] CIGC has

28 Translated from French by the author.

29 The Ministry of Finance is in charge of competition policy.

30 While the government would not have sufficient information to determine precisely the competitive quantity in each period, our results suggest that the monopoly quantity would be significantly lower than the competitive one. Therefore, it is unlikely that the industry could have obtained any government agreement on a quantity close to the monopoly quantity. Using the linear demand specification and setting exogenous factors at their sample mean, we calculated the average quarterly competitive quantity to be $Q_C^c \approx 8{,}789$ tons, and the monopoly quantity $Q_C^m \approx 4{,}545$ tons.

also faced the credible threat that individual producers might bring an antitrust case to the *Conseil*, as occurred for the 1998 ruling, and more recently regarding proposed changes in the production requirements for Comté cheese (Conseil de la Concurrence, 2007).

One element that our result fails to explain is the notable attractiveness of Comté production for dairy farmers located in the eligible region. The rate of satisfaction of new potential entrants (i.e. the ratio of newly attributed production rights to the total demand for new production rights each year) is about 10 per cent, suggesting that the production of Comté milk is more profitable for milk producers than that of generic milk (Colinet *et al.*, 2006). One explanation to this fact is that the share of the milk quota rent that accrues to dairy farmers could be larger in the Comté sector than in other sectors. This is plausible given the existence of purchasing contracts between dairy cooperatives (still mostly owned by milk producers) and ripening facilities, which are negotiated through CIGC. Such industry-wide contracts have the ability to mitigate the bargaining power of ripeners and therefore to reallocate a larger share of the quota rent to milk producers than elsewhere in the dairy sector, where processors are free to exercise their bargaining power. In such a scenario, it becomes necessary for CIGC to limit the entry of new milk producers into Comté production, in order to maintain the quota rent at a level comparable to that in other sectors. Excessive entry into the Comté market would mechanically result in a decrease in the price of ripened Comté cheese, which would erode the size of the milk quota rent available to the industry as a whole and to milk producers in particular.

Appendix

A.1. Data treatment

For Comté cheese, the quarterly series were constructed from monthly data. The quarterly price was constructed as the average of monthly wholesale prices weighted by the quantity sold. A break in the original price series occurs between December 1989 and January 1990, because the method used for reporting sales was modified by CIGC at that time. More specifically, starting from January 1990, sellers of ripened cheese had to report producer prices that included the cost of transporting cheese wheels to the buyer's location. This was not the case previously, since in the early 1980s the majority of transfers from seller to buyer occurred at the factory gate. To account for this change in the reporting methodology, the monthly price series was adjusted so that the new price gap between December 1989 and January 1990 matched the average of the corresponding price gaps in the two preceding years, that is, +1 per cent. The adjustment consisted of inflating all prices from the period January 1985 to December 1989 by 9.52 per cent, to construct a price series that included transportation costs.[31] This seems reasonable since the method used by CIGC during the 1990s to retrieve the producer price net of transportation cost was to deflate all prices by a factor of 10 per cent.[32]

For Emmental cheese, we used the wholesale price of Emmental Est-Central rather than that of generic Emmental, because Emmental Est-Central was the main variety produced in the Comté eligible region during the period.[33] The share of Emmental cheese sold in grated form increased from 18 per cent in 1985 to 57 per cent in 2004. Grated Emmental cheese is mainly used as an ingredient in cooking and is not a good substitute for Comté cheese, which is traditionally consumed by itself (Comté cheese was never sold in grated form during the study period). Therefore, only non-grated Emmental cheese was included in the calculation of the average price of Emmental. A price gap due to a change in methodology occurs between December 1989 and January 1990. The series was corrected by the same method as for the Comté price series. This amounted to inflating prices from January 1985 to December 1989 by 17.5 per cent.[34] A second, minor adjustment had to be made to account for a change in methodology that occurred in January 2001.

Finally, the population variable, used to normalise consumption and income, did not come at a quarterly frequency. Therefore, we replicated the yearly observation in each quarter.

A.2. Derivative of the function ϕ

The function ϕ is defined as

$$\phi(Q_C) = C_1\left(Q_C, k_E\left(\bar{M} - \frac{Q_C}{k_C}\right)\right) - \frac{k_E}{k_C} C_2\left(Q_C, k_E\left(\bar{M} - \frac{Q_C}{k_C}\right)\right).$$

31 This choice was made because there are more observations after the break than before, so that fewer observations had to be adjusted.

32 This latter price series was not available. Furthermore, the 10 per cent deflation was based on a cost estimation made by ripening facilities. Since the deflated price was then used to calculate the contractual price of unripe cheese, there was an incentive on the part of ripening facilities to overstate the cost of transportation in order to pull down the price of unripe cheese.

33 The statistical series for Emmental Est-Central and Emmental were merged by the *Syndicat Interprofessionnel du Gruyère Français* in January 1990. By that date, both varieties had become almost perfect substitutes in consumption and there was no distinguishable price gap between the two. The price series is for non-grated Emmental cheese.

34 The difference between the inflating factor for Comté and Emmental could originate in the price difference between these cheeses (transportation cost represents a larger share of the price of Emmental cheese because this cheese is cheaper) and/or in the fact that Emmental cheese was shipped to more and farther places than Comté cheese, thus leading to larger transportation costs.

To prove that ϕ is a non-decreasing function of Q_C, we note that

$$\phi'(Q_C) = C_{11} - \frac{k_E}{k_C}C_{12} - \frac{k_E}{k_C}C_{21} + \left(\frac{k_E}{k_C}\right)^2 C_{22}$$
$$= \begin{pmatrix} 1 \\ -\frac{k_E}{k_C} \end{pmatrix}' \begin{pmatrix} C_{11} & C_{12} \\ C_{21} & C_{22} \end{pmatrix} \begin{pmatrix} 1 \\ -\frac{k_E}{k_C} \end{pmatrix} \geq 0,$$

where the last inequality follows from the convexity of C.

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Corresponding author: Pierre R. Mérel, Department of Agricultural and Resource Economics, University of California, 1 Shields Avenue, Davis CA 95616, USA. E-mail: merel@primal.ucdavis.edu

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Trade liberalisation, agricultural productivity and poverty in the Mediterranean region

Nadia Belhaj Hassine

Faculty of Economics and Management of Nabeul, Tunisia

Magda Kandil

International Monetary Fund, Washington, DC, USA

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Summary

A widely held view in the economic literature is that productivity growth is an important pathway through which trade liberalisation may alleviate poverty. This paper explores the link between trade openness, agricultural productivity growth and poverty reduction in a panel of Mediterranean countries. Technical efficiency scores and total factor productivity indexes are computed using the latent class stochastic frontier model to account for cross-country heterogeneity in farming production technologies. The relevance of agricultural productivity gains for poverty reduction is investigated through joint estimation of real per capita GDP growth and inequality changes in a dynamic panel setting. The findings illustrate the positive effects of openness on farming efficiency and productivity and give strong support to the view that agricultural productivity growth is a channel for poverty alleviation.

Keywords: openness, agriculture, productivity, poverty, latent class model

JEL classification: C24, C33, D24, F43, Q17

1. Introduction

The impact of international trade on economic growth and poverty is a central issue in the debate surrounding globalisation. Despite the controversy about the causal link between trade openness and economic performance in the literature, the virtues of trade's contributions to faster growth and poverty alleviation are generally recognised.[1] More recently, the Uruguay Round commitments and the ongoing Doha Round of agricultural trade negotiations

*Review coordinated by Paolo Sckokai.

1 See Winters *et al.* (2004) and Nissanke and Thorbecke (2006) for a survey of recent evidences.

refocused attention on the economic linkages between trade, agricultural development and poverty and gave a new impetus to the re-examination of the actual contribution of agricultural growth to poverty reduction. The emphasis given to agricultural development goes far beyond its direct benefit to rural poor livelihoods. Agriculture is convincingly argued to have considerable potential for stimulating broader economic growth through its strong linkages to the wider economy and to offer one of the most promising avenues to poverty alleviation.

The empirical evidence provided by Thirtle *et al.* (2003) illustrates the significant positive contribution of research-led agricultural productivity growth to poverty reduction in Africa and Asia. Ravallion and Chen (2007) emphasise the fundamental role of agricultural development in explaining China's progress against poverty and show that agriculture and rural growth are far more poverty alleviating than non-farm activities. In a cross-country study of the importance of agricultural productivity for economic development, Self and Grabowski (2007) find that agricultural technology plays a crucial role in generating growth and in improving well being. These analyses and others support the beneficial role of agricultural development for the poor.[2] However, although many recognise the potential for trade to generate agricultural productivity gains, they confine their empirical investigations to the links between agriculture, growth and poverty.

International trade is presumed to foster productivity growth through the transfer of technology from more advanced countries, which would confer strong pro-poor benefits on recipient developing economies (Winters, 2002; Cline, 2004; Bardhan, 2006). The productivity-enhancing effects of trade have been widely documented in both macro and case studies, although mainly focusing on manufacturing and industry, with very little specific work on agriculture.[3] Despite the importance of productivity in non-farm activities to sustain poverty alleviation, agricultural productivity would be of most direct interest to the majority of the poor.

The potential of trade and agricultural productivity growth to alleviate poverty will depend upon both the magnitude of overall economic growth and the change in income distribution. The dynamic productivity effects of trade liberalisation can bring about distributional changes, and concerns have been raised about the effects of adverse distributional outcomes that may offset the positive poverty alleviation implications. Although the literature contributes to advancing our understanding of the complex relationships among trade openness, growth, inequality and poverty, it is difficult at present to make convincing generalisations about how productivity and international trade might interact in their effects on growth, inequality and poverty (Nissanke and Thorbecke, 2006; Harrison and McMillan, 2007). The benefits

2 See, for example, Datt and Ravallion (1998), Irz *et al.* (2001), Bravo-Ortega and Lederman (2005), Christiaensen *et al.* (2006).

3 Evidence of a positive stimulus from exports to productivity in the industrial sector may be found in Bigsten *et al.* (2004) and Biesebroeck (2005).

of the trade-induced agricultural growth for poverty reduction remain worthy of further investigation.

The purpose of this paper is to investigate whether agricultural productivity is an important pathway through which trade liberalisation affects the poor. To address this question, we use panel data for 14 Mediterranean countries for 1990–2005 to explore the links between: (i) trade openness and agricultural productivity change, and (ii) farming performance, economic growth and inequality. The countries under investigation have taken steps towards greater integration in the global economy. Indeed, their profile is commensurate with the paper's objectives in many respects. First, these countries are about to start implementing a new agreement on trade in agricultural products under the EU–Mediterranean partnership and the Doha Round of the WTO agreement on agriculture. Second, agriculture is a major sector in these countries, although highly distorted because of trade barriers and protective policies.

As Mediterranean countries go ahead with liberalisation within the framework of the Barcelona Agreement, speculations have arisen regarding the impact of liberalisation in accelerating agricultural development via technology transfer, as well as on the potential impact of this higher agricultural productivity in enhancing growth and reducing poverty.

To shed some light on these issues, the paper's analysis embraces two main objectives: (i) estimating technical efficiency (TE) and total factor productivity (TFP) growth in the Mediterranean agricultural sector, examining the role of international trade in stimulating this growth and (ii) investigating the simultaneous influence of agricultural productivity on income growth and income distribution in order to analyse its impact on poverty. The paper uses appropriate econometric techniques to address these two objectives.

The structure of the paper is as follows: Section 2 presents the methodology for the empirical investigation, Section 3 reviews the data and Section 4 reports the empirical results. Finally, Section 5 synthesises the main findings and draws some conclusions.

2. Methodology

2.1. Efficiency, productivity and the role of trade

The evolving concerns regarding food security, natural resources, rural employment and poverty have sparked a growing literature on agricultural productivity and efficiency. Among the several alternative approaches to estimating agricultural efficiency and multifactor productivity, stochastic frontier models have become very popular.[4] In these models, all producers are assumed to use the same technology. However, farmers who operate in different regions, under various agro-climatic conditions and resource endowments, might not share the same production possibilities. Ignoring the technological

4 See, for example, Battese and Coelli (1995) and Fulginiti *et al.* (2004).

differences in the stochastic frontier model may result in biased efficiency estimates, as unmeasured heterogeneity might be confounded with producer-specific inefficiency (Orea and Kumbhakar, 2004).

The model developed by Greene (2005), and applied by Abdulai and Tietje (2007), considers heterogeneity in the production process through the inclusion of firm-specific random effects. This approach helps to scrutinise inefficiency and heterogeneity. One limitation, however, is in treating heterogeneity as an additional error component, thereby ignoring variations in the parameters of the core production frontier capturing technological differences across producers (Tsionas, 2002).

The random parameter models and the more recently proposed latent class models offer more general approaches to modelling technological heterogeneity (Tsionas, 2002; Orea and Kumbhakar, 2004; Greene, 2005). In the random parameters formulation, heterogeneity takes the form of continuous parameter variation across producers. Besides the practical complications involved in estimating models of this kind and in analysing efficiency, allowing each firm to produce with different technology parameters might not appropriately reflect the fact that some producers in the sample may share the same underlying technology and that they could be efficient relative to their own-group frontier but inefficient compared with other groups using different technologies (Tsionas, 2002).

An alternative formulation that should be better suited to modelling cross-country technological heterogeneity in panel data is the latent class stochastic frontier model (LCSFM).[5] This approach combines the stochastic frontier model with a latent sorting of individuals into discrete groups. Countries within a specific group are assumed to share the same production technology, but this is allowed to differ between groups. Heterogeneity across countries is accommodated through the simultaneous estimation of the probability for class membership and a mixture of several technologies.[6]

The latent class framework assumes the simultaneous coexistence of J different production technologies. Countries in the sample are grouped into different clusters, each using one of these technologies. The number of groups and the class membership are a priori unknown to the analyst. The technology for the jth group is specified as:

$$\ln(y_{it}) = \ln f(x_{it}, t, \beta_j) + v_{it}|_j - u_{it}|_j \quad j = 1, \ldots, J; \quad i = 1, \ldots, N; \quad t = 1, \ldots, T, \tag{1}$$

where subscript i indexes countries, t indicates time, β_j is the vector of parameters for group j and y_{it} and x_{it} are, respectively, the output level and the

5 According to Greene (2005), the latent class model can be considered as an approximation to the random parameters model, in which heterogeneity is modelled as generated by a discrete distribution. He carried out a comparison of both methods, but failed to conclude which model is preferable.

6 For details, see Orea and Kumbhakar (2004) and Greene (2005).

vector of inputs. The error component $v_{it}|_j$ is assumed to be independently and identically distributed (iid) $N(0, \sigma_{v_j}^2)$ and is also independently distributed with respect to the non-negative inefficiency component $u_{it}|_j$.[7]

We adopt the scaled specification for $u_{it}|_j$ by writing it as:

$$u_{it}|_j = \exp\left(z'_{it}\delta_j\right)\, e_{it}|_j, \tag{2}$$

where z_{it} is a vector of country-specific control variables associated with inefficiencies and δ_j is a vector of parameters to be estimated. As common in the literature, the random variable $e_{it}|_j$ is assumed to follow an iid half-normal distribution.[8]

The unconditional likelihood for country i is constructed as a weighted average of its J likelihood functions, with the probabilities of class membership used as weights. The overall log-likelihood function can then be written as:

$$\ln \mathrm{LF} = \sum_{i=1}^{N} \ln\left\{\sum_{j=1}^{J} P_{ij} \prod_{t=1}^{T} \mathrm{LF}_{ijt}\right\}, \tag{3}$$

where LF_{ijt} is the group-specific conditional likelihood function for country i at time t, $\prod_{t=1}^{T} \mathrm{LF}_{ijt}{=}\mathrm{LF}_{ij}$ represents the contribution of country i to the conditional likelihood and P_{ij} is the prior probability of class membership.

An attractive feature of the latent class model is that country-specific probabilities of belonging to a group j can be obtained by conditioning these membership probabilities on a set of country characteristics. These class probabilities can be parameterised as a multinomial logit form:

$$P_{ij} = \frac{\exp(\lambda'_j q_i)}{\sum_j \exp(\lambda'_j q_i)}, \qquad j = 1, \ldots, J, \quad \lambda_1 = 0, \tag{4}$$

where q_i is a vector of country-specific and time-invariant variables that explain probabilities and λ_j is the associated group-specific vector of parameters. λ_1 is normalised to zero to ensure identification of the model.

Maximum likelihood parameter estimates of this mixture of frontier functions can be obtained using the expectation maximisation (EM) algorithm[9] (Caudill, 2003; Greene, 2005). Conventional gradient methods may also be used.

7 Most applications of the LCSFM specify the inefficiency component as iid half-normal (Caudill, 2003; Greene, 2005). We follow Orea and Kumbhakar (2004) in allowing some systematic variation in the inefficiency function.

8 $e_{it|j} \sim N^{+}(0, \sigma_{e_j}^2)$, where ' + ' refers to truncation on the left of zero.

9 EM is an iterative approach where each iteration is made up of two steps: the expectation (E) step, which involves obtaining the expectation of the log-likelihood conditioned over the unobserved data, and the maximisation (M) step, which involves maximising the resulting conditional log-likelihood for the complete data set (Greene, 2001).

Using the estimated parameters and Bayes' theorem, we compute the conditional posterior class probabilities from:

$$P_{j|i} = \frac{\mathrm{LF}_{ij} P_{ij}}{\sum_j \mathrm{LF}_{ij} P_{ij}}. \tag{5}$$

This procedure classifies the sample into J different groups and assigns each country to a group on the basis of the highest posterior probability.[10] Each country's efficiency score can be determined relative to the frontier of the group to which that country belongs. However, this approach ignores uncertainty about the true partitioning in the sample. This somewhat arbitrary selection of the reference frontier can be avoided by evaluating the weighted average efficiency score as follows:

$$\ln \mathrm{TE}_{it} = \sum_{j=1}^{J} P_{j|i} \ln \mathrm{TE}_{it}(j), \tag{6}$$

where $\mathrm{TE}_{it}\ (j) = \exp\ (-u_{it}|_j)$ is the technical efficiency of country i using the technology of class j as the reference frontier.

From the estimated LCSFM, we obtain technical efficiency measures for each country. We now turn to evaluating agricultural TFP growth and to examining the contribution of international trade to productivity improvements. Productivity growth is composed of technological change (TC), technical efficiency change (TEC) and scale economies (SE). Consequently, TFP growth can be computed from:[11]

$$\mathrm{GTFP} = \mathrm{TC} + \mathrm{TEC} + \mathrm{SE} = \frac{\partial \ln f}{\partial t} + \frac{-\partial u_{it|j}}{\partial t} + \frac{(\varepsilon_j - 1)}{\varepsilon_j} \sum_{k=1}^{K} \varepsilon_{kj} \frac{\partial \ln x_{ik}}{\partial t}, \tag{7}$$

where GTFP is agricultural TFP growth, ε_{kj} is the elasticity of output for group j with respect to the kth input and ε_j is the sum of all these input elasticities for a given group j.[12]

TFP growth is explicitly related to the change in efficiency, which is assumed to be a function of a set of variables including international trade. Thus, the parameter estimates of the efficiency term in equation (2) can be

10 The number of classes J is assumed to be given. As estimation with a smaller or a larger number of classes than the true J may result in biased estimates, information criteria such as the Schwarz Bayesian information criteria (SBIC) or the Akaike information criteria (AIC) can be used to determine the appropriate number of classes. These statistics are given by: $\mathrm{SBIC}(J) = -2\mathrm{LF}(J) + K(\mathrm{J})\ln(n)$ and $\mathrm{AIC}(J) = -2\mathrm{LF}(J) + 2K(\mathrm{J})$. $\mathrm{LF}(J)$ is the value of the likelihood function with J classes, $K(J)$ is the number of independent parameters to be estimated and n is the number of observations. The decision rule is to take the model with the lowest AIC or SBIC.

11 See Kumbhakar and Lovell (2000) for the tri-partite decomposition of productivity growth.

12 Since input elasticities vary across groups, productivity change estimates are group-specific. Unconditional productivity measures in equation (7) are obtained as the weighted average of all the TFP components.

used to evaluate the contribution of trade openness to technological catch up (efficiency improvement) and agricultural productivity change.

2.2. Agricultural productivity, growth and poverty

Our analysis of the contribution of agriculture to poverty reduction follows the framework established in Lopez (2004). It takes into account the simultaneous influence of agricultural productivity on growth and income distribution in a dynamic panel setting, inferring their combined effects on poverty.

Following the common practice, poverty measures are characterised in terms of a poverty line, which is defined with respect to the average per capita income and the extent of income inequality (Lopez, 2004; Lopez and Servén, 2006). In this paper, we compute three common poverty measures: the headcount index (Pov_0), the poverty gap index (Pov_1) and the squared poverty gap index (Pov_2). The headcount index (or the 'incidence of poverty') gives the proportion of the population with income below the poverty line. The poverty gap index (or the 'intensity of poverty') indicates how far below the poverty line the poor are. The squared poverty gap (or the 'severity of poverty') captures the inequality among the poor. Assuming a log-normal distribution of income, these poverty measures can be expressed as:[13]

$$
\begin{aligned}
\text{Pov}_0 &= \Phi\left(\frac{\ln(L/Y^R)}{\sigma} + \frac{\sigma}{2}\right),\\
\text{Pov}_1 &= \Phi\left(\frac{\ln(L/Y^R)}{\sigma} + \frac{\sigma}{2}\right) - \frac{Y^R}{L}\Phi\left(\frac{\ln(L/Y^R)}{\sigma} - \frac{\sigma}{2}\right),\\
\text{Pov}_2 &= \Phi\left(\frac{\ln(L/Y^R)}{\sigma} + \frac{\sigma}{2}\right) - 2\frac{Y^R}{L}\Phi\left(\frac{\ln(L/Y^R)}{\sigma} - \frac{\sigma}{2}\right) + \left(\frac{Y^R}{L}\right)^2\\
&\quad e^{\sigma^2}\Phi\left(\frac{\ln(L/Y^R)}{\sigma} - \frac{3\sigma}{2}\right),
\end{aligned}
\tag{8}
$$

where L is the poverty line, Y^R is the average per capita income, Φ is the cumulative normal distribution and σ is the standard deviation of the log-normal distribution, given by $\sigma = \sqrt{2}\ \Phi^{-1}\ ((G+1)/2)$, with G being the income Gini index.[14]

On the basis of these poverty measures, the impact of agricultural productivity growth on poverty can be decomposed into a growth component

13 See Lopez and Serven (2006) for details.

14 The poverty measures presented here belong to the Foster–Greer–Thorbecke class (Foster *et al.*, 1984) given by $\text{Pov}_\theta = \int_0^L [(L-r)/L]^\theta h(r)\mathrm{d}r$, where θ is a parameter of inequality aversion, r is income and $h(\cdot)$ is the density function of income.

and an inequality component and may be expressed as:

$$\frac{\partial \ln \text{Pov}}{\partial \text{GTFP}} = \eta \frac{\partial \ln Y^R}{\partial \text{GTFP}} + \kappa \frac{\partial \ln G}{\partial \text{GTFP}} = \left(\frac{\partial \text{Pov}}{\partial Y^R} \frac{Y^R}{\text{Pov}} \right) \frac{\partial \ln Y^R}{\partial \text{GTFP}} + \left(\frac{\partial \text{Pov}}{\partial G} \frac{G}{\text{Pov}} \right) \frac{\partial \ln G}{\partial \text{GTFP}}, \tag{9}$$

where η and κ denote, respectively, the income growth elasticity of poverty and the inequality elasticity of poverty.

Following Lopez (2004), we estimate the contribution of agricultural productivity to both growth and inequality using the following dynamic simultaneous model:

$$\ln Y_{it}^R - \ln Y_{it-1}^R = \psi \ln Y_{it-1}^R + \tau' W_{it} + \gamma \ln G_{it} + \rho_i + \xi_t + w_{it}^y, \tag{10}$$

$$\ln G_{it} - \ln G_{it-1} = \alpha \ln G_{it-1} + \phi' W_{it} + \chi \ln Y_{it-1}^R + \vartheta_i + \varsigma_t + w_{it}^g, \tag{11}$$

where W represents the set of explanatory variables including agricultural productivity growth, ρ and ϑ are unobserved country-specific effects, ξ and ς are time-specific effects and w_{it}^y and w_{it}^g are the error terms.[15]

Equations (10) and (11) can be employed to obtain estimates of how poverty changes would be associated with agricultural productivity growth. The dynamic structure of the system differentiates between the short- and long-run impacts of agricultural productivity on growth, inequality and poverty. The contribution of farming productivity to poverty changes in the short-term is:

$$\frac{\partial \ln \text{Pov}}{\partial \text{GTFP}} = (\tau_m + \gamma \phi_m)\eta + \kappa \phi_m, \tag{12}$$

whereas, in the long-run, it is:

$$\frac{\partial \ln \text{Pov}}{\partial \text{GTFP}} = -\frac{(\alpha \tau_m - \gamma \phi_m)}{(\psi \alpha - \gamma \chi)} \eta - \frac{(\psi \phi_m - \tau_m \chi)}{(\psi \alpha - \gamma \chi)} \kappa, \tag{13}$$

where the subscript m refers to the mth component of the vector W, which includes agricultural productivity.

3. Data

Our empirical application is based on country-level panel data referring to nine southern Mediterranean countries (SMC) (Algeria, Egypt, Israel,

15 Econometric analyses relating agricultural productivity to per capita income and to inequality can be found in Dollar and Kraay (2002), Kraay (2006) and Self and Grabowski (2007).

Jordan, Lebanon, Morocco, Syria, Tunisia and Turkey) involved in partnership agreements with the EU and the five largest EU Mediterranean countries (France, Greece, Italy, Portugal and Spain) during 1990–2005. Our data set includes observations on agricultural production and input use, international trade, income distribution and a number of other variables that are frequently associated with agricultural productivity and growth. These variables, whose definitions, sources and descriptive statistics are provided in Tables A1 and A2, are used to estimate the stochastic production function [equation (1)], the parametric function of the inefficiency component [equation (2)], the class probabilities [equation (4)], the productivity growth [equation (7)] and the growth and inequality [equations (10) and (11)].

The stochastic production frontier is estimated using data on production of 36 agricultural commodities belonging to six product categories (fruits, shellfruits, citrus fruits, vegetables, cereals and pulses) and on the corresponding use of five inputs (cropland, irrigation water, fertilisers, labour and machines).[16] The six product categories include the main produced and traded commodities in the Mediterranean region.

The inefficiency effect model and the productivity growth equation incorporate an array of control variables representing trade openness, human capital, land holdings, agricultural research effort, land quality and institutional quality.

Three different measures are used to proxy the degree of trade openness of each country: total agricultural trade, agricultural trade barriers and agricultural equipment imports. Agricultural commodities are currently protected with a complex system of ad-valorem tariffs, specific tariffs, tariff quotas and are subject to preferential agreements. The determination of the appropriate level of protection is a fairly complex task. The MAcMaps database constructed by the CEPII provides ad-valorem tariffs, and estimates of ad-valorem equivalent of applied agricultural protection, taking into account trade arrangements (Bouët *et al.*, 2004). Our data on agricultural trade barriers are drawn from this database.[17] The use of agricultural equipment imports as a measure of trade openness is explained by the fact that technology diffusion takes place mainly through capital goods. Accordingly, the productivity effects of openness might then be suitably captured by this variable.

Human capital is included to capture the impact of labour quality and the ability to absorb advanced technology. Land quality, land fragmentation and the distribution of agricultural holdings are often cited as potential sources of inefficiency in agriculture (Vollrath, 2007), and agricultural research has been shown to contribute to production efficiency and to sustained productivity by developing yield increasing technologies.

Institutional quality includes various institutional variables considered as indicators of a country's governance, namely, political stability, government

16 We construct aggregate output and input indexes for each product category using the Tornqvist and Eltetö–Köves–Szulc (EKS) indexes. See Eltetö and Köves (1964) and Szulc (1964).

17 Available at http://wits.worldbank.org/witsweb/default.aspx.

effectiveness and control of corruption. These variables reflect the ability of the government to provide sound macroeconomic policies and impartial authority to protect property rights and enforce contracts. Improved institutional quality is thought to enhance farming efficiency and productivity, as it may facilitate human capital accumulation, appropriate technology adoption and provision of rural infrastructure as well as it may boost the incentives to produce and invest (Self and Grabowski, 2007; Vollrath, 2007).

As determinants of the latent class probabilities, we consider country averages of five separating variables: total agricultural machinery, total applied fertilisers, agricultural land, average farm size and rainfall levels. Machinery and fertilisers help to identify countries endowed with modern inputs. Average farm size captures the differences in the scale of agricultural holdings across countries and distinguishes countries with an important proportion of small farms (Vollrath, 2007). Agricultural land and rainfall levels capture the influence of resources endowments and climatic conditions on class membership.

Finally, to estimate the growth and inequality equations, we use per capita GDP in 2000 international dollars as a measure of mean income and the income Gini index as a proxy for inequality. The data on income distribution are drawn from Dollar and Kraay's (2002) and UN-WIDER World Income Inequality Databases. Although these data sets contain a large amount of information about the distribution of income within many countries, they form an unbalanced and irregularly spaced panel of observations. We approximate the missing values using a simple linear time-trend forecast.[18]

For calculating poverty measures, we need to define the poverty line. This arbitrary choice is problematic, since our study includes both developing and developed countries. Our analysis reports the results for the poverty measures and the growth and inequality elasticities of poverty using two specific poverty lines: L1, set at 50 per cent of the mean per capita GDP (in constant 2000 International dollars), and L2, set at the mean income of the three first quintiles in each country.

4. Estimation results

4.1. The efficiency and productivity model

This empirical exercise involves basically a three-step analysis of agricultural productivity performance across Mediterranean countries. First, a Cobb–Douglas parameterisation of the technology frontier is employed and the latent class model of equation (1) is estimated using maximum likelihood via the EM algorithm.[19] Despite its inherent restrictive properties, the Cobb–Douglas functional form has been widely used in the empirical estimation of productivity models since more flexible forms, such as the translog function, consume more degrees of freedom and are generally associated with a serious multi-collinearity problem, which reduces the precision of the

18 A similar approximation has been used in Sala-i-Martin (2002).
19 The estimation procedure was programmed in Stata 9.2.

estimated parameters. Second, efficiency levels and productivity growth are computed for each agricultural commodity group and country. Third, the determinants of agricultural productivity growth are investigated focusing on the role of international trade.

In each country, we carried out estimations at different levels of aggregation, both for each agricultural commodity group and for the whole agricultural sector.

In estimating the latent class model, we begin by examining the class selection issue. The SBIC and AIC test results support the class segmentation of the model and indicate that the model with four classes is preferred for citrus fruits, shell fruits, vegetables and for the pooled model, while the preferred number of classes for the remaining product categories is three. Table 1 presents the results of estimating the input elasticities of the production frontier.[20]

For the production function, we obtain fairly reasonable estimates. The input elasticities are globally positive and significant at the 10 per cent level or lower. The differences of the estimated factor elasticities among classes seem to support the presence of technological differences across countries. Water and cropland have globally the largest elasticity, indicating that the increase of Mediterranean agricultural production depends mainly on these inputs. The role of water is especially important, indicating that Mediterranean crops are highly water intensive and water is the most limiting and precious input in this region. Labour and, to a lesser extent, machines seem also to be important factors in crop production. Fertilisers, although significant, appear to have a limited effect on Mediterranean production. This may be explained by the fact that farmers in some regions tend to use fertilisers as complementary input to organic manure, which is much less expensive.

In addition to production elasticities, the estimated technology frontiers provide a measure of technical change: a positive sign on the time trend variable reflects technical progress. We find significant shifts in the production frontier over time, indicating gains in technical change.

Table 1 reports also the results concerning the determinants of agricultural production efficiency among the selected countries. The estimated coefficients of the efficiency term are statistically significant at conventional levels and have globally the expected signs. Trade openness seems to exert a significant impact on improving efficiency in the Mediterranean farming sector.[21] Educational attainment, land quality, agricultural research effort and institutional factors appear also to contribute to enhancing efficient input use. As expected, the unequal distribution of agricultural land and, to a lesser extent, land fragmentation have significant adverse effects on efficient resource use. Concentrated land distribution is generally associated with underutilisation of resources. A more equitable distribution of land holdings may help to

20 In the interest of space limitation, we describe the results using pooled data. Results of AIC and SBIC as well as estimates for specific crops are available from the authors upon request.

21 The model has been estimated using all the three measures of trade openness. The results obtained with the other two variables (total agricultural trade and agricultural trade barriers) are very similar.

Table 1. Latent class model parameter estimates

	Class 1	Class 2	Class 3	Class 4
Production frontier				
Land	0.309***	0.261***	0.444***	0.216***
Water	0.275***	0.289***	0.276***	0.333***
Labour	0.236***	0.262***	0.141*	0.144**
Fertilisers	0.107*	0.092*	0.127*	0.111*
Machines	0.097*	0.163*	0.136**	0.327***
Time	0.017***	0.064**	0.009**	0.008*
Intercept	0.551**	0.760**	0.022	0.122
Efficiency term				
Land Gini	0.212***	0.169***	0.175***	0.123***
Land fragmentation	0.038**	0.002*	0.058**	0.024*
Land quality	−0.044**	−0.043*	−0.052***	−0.011*
Trade openness[a]	−0.157***	−0.135***	−0.268***	−0.165***
Human capital	−0.095***	−0.098**	−0.156**	−0.149**
R&D	−0.004*	−0.002*	−0.002**	0.001*
Government effectiveness	−0.026	−0.003*	−0.013**	0.003***
$\Gamma = \sigma_e^2/\sigma_s^2$	0.721***	0.829***	0.784***	0.891***
Probabilities				
Fertilisers consumption		−0.073	0.144**	−0.991**
Agricultural machinery		0.079*	−0.034	0.472***
Agricultural land		0.037***	0.045**	0.408***
Average holdings		−0.026**	0.352*	0.093**
Rain		−0.006*	0.013**	0.262**
Intercept		−1.360	−1.359*	−3.291**
Log-likelihood		−274.333		
Number of observations		1,344		

Notes: The variables in the production frontier and efficiency function are in natural logarithm. $\sigma_s^2 = \sigma_e^2 + \sigma_v^2$. The significance at the 10, 5 and 1 per cent levels is indicated by *, ** and ***, respectively. A negative sign in the inefficiency model means that the associated variable has a positive effect on technical efficiency.
[a]Measured as agricultural equipment import.

ensure the equality of marginal products of inputs across farms (Vollrath, 2007). On the other hand, land fragmentation may lead to a sub-optimal usage of factor inputs owing to inadequate monitoring, the inability to use certain types of machines and wasted space along the borders. Inequitable land distributions and high fragmentation of land may also reflect the existence of an important number of small farms with limited financial resources, low skills and inefficient traditional production methods.

The results of the latent class probability functions, also reported in Table 1, show that the coefficients are globally significant, indicating that the variables included in the model provide useful information in classifying the sample.[22] We had no a priori expectation about the sign of these coefficients, as positive values on the separating variables' coefficients in one class indicate that higher values of these variables increase the probability of assigning a country into this class, whereas negative parameters suggest that the probability of class membership decreases with an increase of the corresponding variables. For example, increasing total applied fertilisers increases the probability of a country belonging to class 3. Wider total agricultural areas increase the probability of membership in the three last classes, whereas higher average farm size reduces the probability of belonging to the class 2.

Table 2 summarises the estimated prior and posterior class probabilities as well as the grouping of countries between the different classes in the pooled and specific commodity models. The posterior class probabilities are, on average, very high (70 per cent or more). The classification resulting from these probabilities shows globally that Algeria, Israel, Jordan, Lebanon and Portugal belong to the same group characterised by relatively low agricultural production levels. Further, the sample of countries exhibits a similar pattern of specialisation on the basis of a strong presence of fruits and vegetables, limited use of modern technical inputs, relatively poor agricultural land endowment, small average farm size and low precipitation. The second group of countries comprises Greece, Morocco, Syria and Tunisia, characterised by higher production levels and higher precipitation. The remaining country groups include Egypt, France, Italy, Spain and Turkey. The average production level of these countries is significantly larger than that in other classes; the use of modern technical input, the endowment in agricultural land, the average farm size and the precipitation level are higher than that of the other groups. These countries show a common cropping pattern in which cereal crops account for an important share.

The average efficiency scores, TFP growth and the contribution of international trade to productivity changes, estimated using equations (6) and (7), respectively, are reported in Table 3. The results show that productivity has increased in all the Mediterranean countries, with SMC registering relatively higher average rates of productivity gain than EU countries. Significant differences in technical efficiency performance are apparent among commodity groups and countries. On average, over the period under consideration, EU countries exhibited better efficiency levels than SMC.

The parameter estimates of the efficiency component in Table 1 show a significant positive impact of trade openness on technological catch-up. We use these estimates to measure the effect of trade, through changes in efficiency, on agricultural productivity growth. The results in Table 3 show a positive contribution of trade openness to improvements in agricultural productivity in the Mediterranean region. The findings show that international trading

22 The impact of the different variables on the class probabilities could also be measured by the marginal effects. We focus only on the estimated parameters since their sign depends on the sign of the marginal effects.

Table 2. Prior and posterior probabilities

Class	Countries	Prior probabilities	Posterior probabilities
Fruits			
1	Algeria, Israel, Lebanon, Jordan	0.625	0.793
2	Egypt, Morocco, Portugal, Syria, Tunisia, Greece	0.584	0.875
3	Spain, France, Italy, Turkey	0.700	0.788
Citrus			
1	Algeria, Portugal, Jordan, Lebanon	0.609	0.762
2	Greece, Morocco, Syria, Israel, Tunisia	0.720	0.956
3	France, Spain, Egypt	0.567	0.759
4	Italy, Turkey	0.706	0.873
Shell fruits			
1	Algeria, Jordan, Lebanon, Egypt	0.471	0.724
2	Morroco, Portugal, Tunisia, Israel	0.609	0.762
3	France, Greece, Syria	0.667	0.842
4	Italy, Spain, Turkey	0.600	0.990
Vegetables			
1	Israel, Jordan, Lebanon	0.635	0.805
2	Portugal, Algeria, Morocco, Syria, Tunisia	0.627	0.787
3	France, Egypt, Greece	0.591	0.752
4	Spain, Turkey, Italy	0.850	0.864
Cereals			
1	Algeria, Israel, Jordan, Lebanon	0.513	0.784
2	Portugal, Tunisia, Greece, Morocco, Syria, Spain	0.610	0.791
3	Egypt, France, Italy, Turkey	0.741	0.990
Pulses			
1	Portugal, Israel, Jordan, Lebanon	0.633	0.749
2	Algeria, Greece, Italy, Morocco, Syria, Tunisia	0.645	0.781
3	France, Spain, Egypt, Turkey	0.702	0.781
Total pool			
1	Algeria, Portugal, Israel, Jordan, Lebanon	0.674	0.712
2	Morocco, Greece, Syria, Tunisia	0.559	0.796
3	Spain, France, Egypt	0.634	0.826
4	Italy, Turkey	0.625	0.835

Note: Prior probability and posterior probability represent, respectively, the averages over the countries of the estimated prior and posterior probabilities for the highest probability classes.

Table 3. Efficiency scores and TFP index growth

	Fruits			Citrus			Shell			Vegetables			Cereals			Pulses			Pool		
	TE[a]	GTFP[b]	TO[c]	TE	GTFP	TO	TE	GTFP	TO	TE	GTFP	TO	TE	GTFP	TO	TE	GTFP	TO	TE	GTFP	TO
Algeria	0.54	2.88	0.44	0.42	2.39	0.80	0.60	−1.19	0.05	0.68	0.62	0.75	0.55	1.78	0.25	0.64	−0.58	0.14	0.60	1.14	0.40
Egypt	0.58	1.37	0.17	0.66	1.64	0.19	0.59	−0.90	0.08	0.44	4.90	0.34	0.58	−0.14	0.07	0.59	1.61	0.05	0.60	1.16	0.15
France	0.92	1.08	0.12	0.83	−1.18	0.09	0.96	0.60	0.16	0.99	0.55	0.07	0.99	1.21	0.40	0.98	1.09	0.06	0.98	0.96	0.13
Greece	0.63	1.47	0.18	0.71	1.73	0.28	0.63	−1.65	0.43	0.65	−0.85	0.35	0.66	1.91	0.38	0.68	1.03	0.07	0.68	0.85	0.28
Israel	0.68	1.54	0.04	0.79	1.19	0.13	0.67	1.74	0.25	0.71	2.13	0.32	0.48	−0.74	0.04	0.64	2.74	0.05	0.67	1.82	0.14
Italy	0.89	1.51	0.13	0.75	1.55	0.96	0.71	0.74	0.75	0.81	1.41	0.41	0.74	1.79	0.38	0.79	1.10	0.05	0.81	1.45	0.45
Jordan	0.61	0.97	0.75	0.67	1.22	0.56	0.63	1.74	0.38	0.79	1.66	1.07	0.35	−0.89	0.87	0.65	1.72	0.20	0.66	1.34	0.63
Lebanon	0.88	1.31	0.21	0.77	1.28	0.78	0.87	1.62	0.29	0.82	1.95	0.76	0.61	1.98	0.51	0.81	−0.47	0.09	0.79	1.61	0.44
Morocco	0.62	−0.46	0.57	0.86	1.12	0.82	0.67	2.94	1.78	0.77	1.45	0.65	0.63	−0.25	2.50	0.63	1.32	0.34	0.74	1.05	1.11
Portugal	0.53	0.38	0.22	0.63	1.39	0.08	0.51	0.24	0.89	0.71	−0.41	0.09	0.64	1.92	0.03	0.56	−0.25	0.10	0.61	0.79	0.26
Spain	0.79	1.59	0.56	0.85	1.01	0.98	0.68	−2.37	2.26	0.88	1.78	0.64	0.76	1.63	0.71	0.69	0.73	0.29	0.80	0.96	1.04
Syria	0.65	1.33	0.74	0.79	0.99	0.79	0.70	3.04	0.50	0.74	2.45	0.74	0.77	2.76	1.16	0.76	1.42	0.37	0.74	2.01	0.79
Tunisia	0.64	0.74	0.05	0.64	1.03	0.43	0.69	0.31	0.10	0.73	1.62	0.05	0.68	0.93	0.30	0.65	1.58	0.01	0.66	1.07	0.16
Turkey	0.88	1.79	0.26	0.88	2.19	1.19	0.88	2.08	0.52	0.82	1.87	1.13	0.85	1.89	0.82	0.79	2.26	0.26	0.83	2.08	0.70

[a]Technical efficiency score (the index varies between 0 and 1).
[b]TFP growth (per cent).
[c]Contribution of trade openness (measured by Agricultural equipment import) to agricultural TFP growth (per cent).

opportunities should generate larger benefits in SMC such as Morocco, Syria, Turkey and Jordan.

4.2. The poverty–agricultural productivity nexus

Assessing the role of agricultural productivity gains in alleviating poverty requires the empirical evaluation of the impact of productivity on both growth and distributional change, as well as measuring the elasticity of poverty with respect to each of them. We first compute the growth and inequality elasticities of poverty, for two different poverty lines, using the three poverty measures mentioned before. Second, we estimate equations (10) and (11) to investigate the impact of agricultural TFP growth on income and inequality. Finally, we analyse the short- and long-run poverty outcomes using equations (12) and (13).

The estimated dynamic equations in (10) and (11) are potentially biased by endogeneity arising from correlation between the explanatory variables and unobserved country-specific effects. To deal with this endogeneity problem, we use the generalised method of moments (GMM) approach, as suggested by Blundell and Bond (1998).[23]

The estimation results of the equations for real per capita GDP growth and inequality are presented in Table 4. We report the results obtained using three different measures of trade openness, namely agricultural equipment import (column 1), total agricultural trade (column 2) and agricultural trade barriers (column 3).

The consistency of the system GMM estimator is checked using two specification tests. The first tests for serial correlation and the second addresses the instrument validity issue using the Sargan test of overidentifying restrictions. On the basis of the results of these tests, the assumptions of no second-order serial correlation and of the validity of the instrument set are not rejected.

A number of significant results emerge from the empirical analysis. First, the findings indicate that an increase in inequality would positively affect growth, whereas an increase in income should reduce inequality. The point estimates of the coefficients for these variables suggest, however, small potential impacts.

Agricultural productivity growth appears as a key factor to economic growth and income distribution in Mediterranean countries. This is evident by the positive and statistically significant coefficient of GTFP in the per capita real GDP growth equation and the significant negative coefficient in the equation explaining inequality.

Trade openness seems as well to be positively related to growth and negatively associated with inequality. Our results suggest that increased trade and

23 This method involves estimating simultaneously a two-equation system, consisting of the differenced equation and the original level equation, subject to appropriate cross-equation restrictions that constrain the coefficient vectors in the level and differenced equations to be identical. This approach uses lagged differences as instruments for contemporaneous levels, in addition to the lagged levels as instruments for contemporaneous differences.

Table 4. Determinants of growth and inequality

Variables	Equipment import	Total trade	Trade barriers
Income growth equation			
Lagged income	−0.0481*	−0.0205***	−0.0224**
Income Gini	0.0089*	0.0042*	0.0046*
GTFP	0.0793***	0.0756***	0.0689***
International trade	0.0326***	0.0221*	−0.0241***
Human capital	0.0485***	0.0273*	0.0282**
R&D	0.0091**	0.0031*	0.0042*
Land fragmentation	−0.0032*	−0.0012*	−0.0026*
Land Gini	−0.0034***	−0.0045***	−0.0052***
Average holdings	0.0028*	0.0027	0.0023*
Fertilisers consumption	0.0032*	0.0017	0.0026*
Agricultural machinery	0.0044***	0.0053*	0.0068**
Control of corruption	0.0013	0.0016*	0.0018*
Government effectiveness	0.0007***	0.0004***	0.0003***
Political stability	0.0001*	0.0001*	0.0001*
M1: first-order serial correlation	$z = -4.65$	$z = -4.62$	$z = -4.59$
	$p = 0.00$	$p = 0.00$	$p = 0.00$
M2: second-order serial correlation	$z = 1.13$	$z = 0.93$	$z = 1.02$
	$p = 0.26$	$p = 0.35$	$p = 0.31$
Sargan instrumental validity test	$\chi^2(166) = 142$	$\chi^2(144) = 126$	$\chi^2(154) = 135$
	$p = 0.91$	$p = 0.86$	$p = 0.86$

(*Continued on the next page*)

Table 4. (*continued*)

Variables	Equipment import	Total trade	Trade barriers
Inequality change equation			
Lagged income Gini	−0.0592***	−0.0661***	−0.0695***
Lagged income	−0.0326*	−0.0384*	−0.0473*
GTFP	−0.2994***	−0.2241***	−0.1832**
International trade	−0.0343***	−0.0134*	0.0495**
Human capital	−0.0771**	−0.1830***	−0.2290***
R&D	0.0015***	0.0020***	0.0018***
Land fragmentation	0.0079***	0.0092**	0.0095**
Land Gini	0.0084*	0.0038***	0.0044**
Average holdings	−0.0051**	−0.0061*	−0.0094**
Fertilisers consumption	−0.0006***	−0.0006***	−0.0007**
Agricultural machinery	−0.0961***	−0.1181***	−0.1011***
Control of corruption	0.0049*	0.0051	0.0048*
Government effectiveness	−0.0007	0.0003	0.0008
Political stability	−0.0002**	−0.0002*	−0.0002
M1: first-order serial correlation	$z = -6.29$	$z = -6.47$	$z = -6.52$
	$p = 0.00$	$p = 0.00$	$p = 0.00$
M2: second-order serial correlation	$z = 1.10$	$z = 1.05$	$z = 0.90$
	$p = 0.27$	$p = 0.29$	$p = 0.37$
Sargan instrumental validity test	$\chi^2(102) = 84$	$\chi^2(104) = 85$	$\chi^2(98) = 83$
	$p = 0.90$	$p = 0.91$	$p = 0.86$
Number of observations	224	224	224

Notes: The dependent variables are the real per capita GDP growth and the income Gini index change. The significance at the 10, 5 and 1 per cent levels is indicated by *, ** and ***, respectively.

reduced tariff barriers are likely to contribute to faster growth as well as to inequality reduction. On the other hand, we find that research effort, as computed by R&D expenditures, would contribute to growth, but increasing inequality.

Factors such as land fragmentation and inequality of land holdings appear to thwart economic growth and to accentuate income disparity. The point estimates of the coefficients of these variables indicate that increasing average farm sizes and reducing disparities in land ownerships would strongly help to close the income inequality gap. Modern technical inputs appear also to contribute to economic growth and inequality reduction.

The results for human capital indicate that greater endowment of educated workforce would enhance growth, as evident by the positive and statistically significant coefficient in the per capita GDP growth equation. This variable enters the inequality equation with a negative and significant coefficient, indicating that countries with better education systems would be less unequal.

A number of institutional quality measures are taken into consideration. The results are very reassuring regarding the positive effects of institutional quality on real per capita GDP growth. Control of corruption, government effectiveness and political stability have positive and significant effects, indicating an increase in real per capita GDP growth as a response to an improvement in these institutional indicators. Surprisingly, however, improvement in the institutional quality indicators does not appear to have a significant bearing on inequality changes.

Table 5 reports the elasticity of poverty with respect to growth and inequality. A review of these results confirms the common claim that poverty, regardless of its measure, decreases with real growth. The impact of growth on poverty seems stronger in Mediterranean EU countries than in the south side. Countries such as Lebanon, Syria and Turkey, where the growth elasticity is low in absolute value, will find it harder than France, Italy, Spain and Greece to achieve fast poverty reduction. In line with the evidences in the empirical literature, we also find that inequality is positively related to poverty. In general, the evidence is robust. Regardless of the poverty measure and benchmark, the results indicate that income inequality exacerbates and perpetuates poverty.

Poverty seems to react positively to inequality and negatively to growth, and the absolute sizes of these two effects are positively correlated.[24] On average, both growth and inequality have smaller effects on poverty in the SMC. The growth and inequality elasticities of poverty appear also to be larger, in absolute values, for the L2 poverty line.

Overall, our results suggest that policies that support both higher growth and lower inequality would induce poverty reduction. However, pro-growth policies worsening income distribution may have ambiguous poverty

24 It is interesting to note, however, that the distribution of the elasticity coefficient is not robust to the various measures of poverty. The distribution of the squared poverty gap is in sharp contrast to the distribution of the headcount and poverty gap.

Table 5. Growth and inequality elasticities

	Headcount (Pov_0)				Poverty gap (Pov_1)				Squared poverty gap (Pov_2)			
	L1		L2		L1		L2		L1		L2	
	GE[a]	IE[b]	GE	IE	GE	IE	GE	IE	GE	IE	GE	IE
Algeria	−4.20	2.13	−8.52	3.90	−2.74	3.42	−3.42	5.35	−1.55	19.78	−1.88	316.13
Egypt	−3.71	1.95	−6.04	3.30	−2.54	3.22	−3.08	4.71	−1.45	13.47	−1.71	114.51
France	−6.08	2.69	−11.08	4.32	−3.68	4.05	−4.38	5.08	−2.06	54.77	−2.40	455.00
Greece	−4.20	2.12	−7.87	3.76	−2.75	3.41	−3.39	5.20	−1.56	18.54	−1.87	259.39
Israel	−3.44	1.87	−6.77	3.71	−2.35	3.12	−3.03	5.15	−1.34	18.39	−1.67	214.10
Italy	−4.42	2.21	−8.33	3.88	−2.92	3.52	−3.58	5.34	−1.65	14.29	−1.97	246.95
Jordan	−3.38	1.85	−7.42	3.94	−2.27	3.10	−3.00	5.39	−1.30	17.74	−1.65	315.08
Lebanon	−1.77	1.25	−2.98	2.95	−1.09	2.36	−1.51	4.33	−0.63	149.12	−0.84	884.11
Morocco	−3.17	1.77	−6.21	3.68	−2.17	3.01	−2.83	5.12	−1.24	18.16	−1.56	192.68
Portugal	−3.56	1.92	−7.65	4.01	−2.41	3.18	−3.16	5.47	−1.37	12.51	−1.74	286.41
Spain	−4.60	2.25	−8.24	3.65	−2.96	3.55	−3.54	5.09	−1.67	24.65	−1.96	244.23
Syria	−2.59	1.56	−4.42	3.13	−1.75	2.75	−2.26	4.52	−1.00	40.51	−1.25	166.43
Tunisia	−3.02	1.71	−6.02	3.61	−2.03	2.93	−2.66	5.03	−1.16	25.33	−1.47	251.81
Turkey	−2.70	1.60	−5.71	3.62	−1.82	2.80	−2.43	5.04	−1.03	39.46	−1.34	330.83

Notes: L1: poverty line is 50 per cent of mean per capita GDP; L2: poverty line is the mean income of the three first-quintile shares.
[a]Growth elasticity.
[b]Inequality elasticity.

Table 6. Short- and long-run poverty effects of agricultural TFP growth

	Headcount (Pov_0)				Poverty gap (Pov_1)				Squared poverty gap (Pov_2)			
	L1		L2		L1		L2		L1		L2	
	SR	LR	SR	LR	SR	LR	SR	LR	SR	LR	SR	LR
Algeria	−0.96	−8.84	−1.82	−15.66	−1.23	−16.78	−1.86	−26.81	−6.03	−106.29	−94.67	−1713.54
Egypt	−0.87	−8.18	−1.45	−14.00	−1.16	−15.83	−1.64	−23.56	−4.14	−72.13	−34.37	−620.02
France	−1.27	−10.67	−2.14	−16.29	−1.49	−19.59	−1.85	−24.73	−16.53	−295.75	−136.23	−2466.46
Greece	−0.96	−8.79	−1.73	−15.32	−1.23	−16.72	−1.81	−26.02	−5.66	−99.56	−77.70	−1405.77
Israel	−0.82	−7.92	−1.63	−15.76	−1.11	−15.41	−1.77	−25.98	−5.60	−98.89	−64.14	−1160.24
Italy	−1.00	−9.14	−1.80	−15.67	−1.28	−17.21	−1.87	−26.66	−4.40	−76.45	−73.99	−1338.23
Jordan	−0.81	−7.85	−1.75	−16.59	−1.10	−15.35	−1.84	−27.30	−5.40	−95.39	−94.34	−1707.99
Lebanon	−0.51	−5.64	−1.11	−14.08	−0.79	−12.10	−1.41	−22.51	−44.64	−808.45	−264.41	−4795.04
Morocco	−0.77	−7.56	−1.58	−15.96	−1.07	−14.93	−1.75	−25.95	−5.52	−97.70	−57.73	−1044.13
Portugal	−0.85	−8.12	−1.79	−16.82	−1.14	−15.69	−1.88	−27.63	−3.85	−66.97	−85.77	−1552.37
Spain	−1.03	−9.24	−1.72	−14.48	−1.29	−17.35	−1.79	−25.33	−7.50	−132.63	−73.17	−1323.49
Syria	−0.66	−6.79	−1.27	−14.13	−0.96	−13.79	−1.52	−23.06	−12.19	−219.09	−49.86	−901.94
Tunisia	−0.74	−7.33	−1.54	−15.64	−1.03	−14.58	−1.71	−25.57	−7.66	−136.65	−75.40	−1364.92
Turkey	−0.69	−6.94	−1.52	−15.95	−0.98	−14.03	−1.69	−25.77	−11.88	−213.37	−99.02	−1793.62

Note: SR is short-run semi-elasticity of poverty; LR is long-run semi-elasticity of poverty.

outcomes, since the poverty-reducing effects of growth may be outweighed by the poverty-raising effects of inequality.

The findings in Table 4 indicate that several pro-growth policies, such as agricultural trade, agricultural productivity growth, greater endowment of educated labour force and more equitable land distribution would reduce inequality and are, therefore, pro-poor. On the other hand, agricultural R&D expenditures possibly increase inequality and, thus, present a trade-off between their growth and inequality outcomes. Some countries may be willing to tolerate modest deteriorations in income equality in exchange for faster growth. This trade-off is problematic in countries such as Lebanon and Syria, where the small positive growth impact would not be enough to offset the inequality poverty-raising effect.

Table 6 reports the short- and long-run poverty effects of agricultural productivity growth in terms of semi-elasticity of poverty with respect to GTFP, as defined in equations (12) and (13).

As expected, agricultural productivity has an important role to eradicate poverty. Improving farming performance appears to strongly benefit the poor in the long run. Using the headcount measure and the L1 benchmark, poverty decreases in the short run with a semi-elasticity estimate ranging from a low of −0.51 in Lebanon to a high of −1.27 in France, whereas the long-run productivity impact on poverty ranges from −5.64 in Lebanon to −10.67 in France.[25] The results suggest that agricultural growth tends to play a more prominent role in reducing poverty in EU countries. The evidence is robust with respect to the L2 benchmark and appears to be larger in absolute values. The poverty gap and the squared poverty gap seem to be highly sensitive to growth in agricultural productivity, suggesting that improvements in farming practices would help to reduce the depth of poverty and to mitigate inequality.

5. Concluding remarks

Proponents of globalisation identify strong benefits from trade liberalisation in terms of resource allocation, economic growth and poverty alleviation. Despite the controversy that surrounds the trade issues, there is widespread acceptance that relatively open policies contribute significantly to development.

The existing empirical literature has been relatively successful in examining the association between trade openness, growth and poverty; it has, however, much less to say about the specific role of agricultural productivity gains.

The analysis of this paper has focused on the impact of trade liberalisation on agricultural productivity and poverty in the Mediterranean region. Agriculture is a vital sector in the Mediterranean economies, as it represents an

25 A semi-elasticity of −10.67 indicates that an acceleration of agricultural TFP growth from 1.2 to 1.3 per cent would reduce poverty by around 10 per cent in the long run.

important source of income and output and employs a large segment of impoverished population. The critical rural dimension of poverty in the Mediterranean region points to a potential central role for the agricultural sector in poverty eradication.

Agriculture has always been subject to various protection mechanisms that have distorted market incentives and resulted in inefficient allocations of resources. As the Mediterranean region proceeds with more plans for trade liberalisation, attention has focused on its effects on agricultural productivity and poverty reduction.

To that end, our analysis examines the effects of trade openness on agricultural productivity in the Mediterranean basin and assesses how farming performance impinges on poverty through the projected interaction with growth and inequality.

The distinguishing aspect of this study is the attempt to account for heterogeneity in cross-country agricultural production in the estimates of technical efficiency and productivity change. The methodology follows the LCSFMs where output includes 36 agricultural commodities, grouped in six categories, and five input variables. Estimates support the presence of technological differences across countries. Mediterranean crops appear to be highly water intensive, which limits productive capacity given the scarcity of water in the region. Although there are important inefficiencies in Mediterranean agricultural production, the evidence reveals positive rates of productivity growth in most of the selected countries.

One of the salient features of the regression results is that trade openness exerts a significant ameliorating influence on the incidence of poverty in Mediterranean countries. We find evidence that severe trade restrictions may increase income inequality and adversely affect growth. Through this channel, stronger trade liberalisation should have beneficial effects on poverty.

The impact of increased international trade on poverty is further reinforced through the indirect channel. Opening up to foreign trade seems to facilitate catching up with the best practice technology, providing substantial support for the view that openness promotes productivity growth through technology transfers. Agricultural productivity gains would lead to both faster growth and lower income inequality. Therefore, agricultural development seems to have positive effects on the society as a whole but, given high concentration of poverty in rural areas, the poor would benefit more than proportionately. Our findings indicate that farming performance tends to play a more prominent role in reducing poverty in advanced countries. Behind this difference, there is a larger incidence of poverty in developing countries that necessitates a longer path of sustained improvement in agricultural productivity before a significant effect on poverty alleviation is realised.

To summarise, the paper's results support the benefits of trade liberalisation on agricultural growth and poverty reduction in the Mediterranean region. As the agriculture sector is likely to reap the benefits of liberalisation and trade openness, the income inequality gap is likely to shrink and the incidence of poverty is likely to decrease. Such added benefits provide direct testimony that developing countries in the Mediterranean basin should be more actively pursuing efforts to increase trade linkages and integration.

It is necessary to emphasise, however, that the added benefits of trade liberalisation are contingent on complementary efforts that would reinforce the positive effects on per capita income growth and poverty reduction. The paper's evidence provides a menu of complementary structural, policy and institutional measures that should be in place to ensure the maximum added benefits of trade liberalisation on growth and welfare in the Mediterranean economies.

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Appendix

Table A1. Variable definitions and sources of data

Variables	Definitions	Units	Sources
Agricultural land	Total agricultural land	Per cent of land area	WDI
Agricultural machinery	Total wheel and crawler tractors	Machinery/100 ha of arable land	WDI
Average holdings	Average farm size for the commodities included in the analysis	Ha	FAO
Control of corruption	Control among public and private officials, extent of bribery, etc.	Index value[a]	Kaufmann *et al.* (2007)
Equipment import	Agricultural machinery and equipment imports	Per cent of agricultural value added	FAO
Fertilisers consumption	Total fertiliser consumption	100 g/ha of arable land	WDI
Fertilisers	Fertilisers use by commodity	1,000 tons	FAO, FEMISE
Government effectiveness	Efficiency of country's bureaucracy, state's ability to create national infrastructure, etc.	Index value[a]	Kaufmann *et al.* (2007)
Human capital	Average years of schooling in the population over age25	Number of years	Barro and Lee (2000)
Income	Per capita GDP	2000 international dollar	WDI
Income Gini	Income Gini index	Percentage	Dollar and Kraay (2002), UN-WIDER
Labour	Labour use by commodity	Million of days worked	FAO, FEMISE
Land	Land use by commodity	Million hectares	FAO, FEMISE
Land fragmentation	Share of holdings under 5 ha	Per cent of agricultural land	FAO

(*Continued on the next page*)

Table A1. (*continued*)

Variables	Definitions	Units	Sources
Land Gini	Inequality in land distribution measured by the Gini coefficient for land holdings	Percentage	FAO
Land quality	Share of irrigated area	Per cent of agricultural land	WDI
Machines	Wheel and crawler tractors use by commodity	Million hours	FAO, FEMISE
Output	Quantity of agricultural output	Million tons	FAO
Political stability	The unlikelihood of armed conflict, ethnic tensions, terrorist threats, etc.	Index value[a]	Kaufmann *et al.* (2007)
Rain	Average precipitations (1961–2005)	Km^3/year	FAO
R&D	Public and private agricultural R&D expenditures	Million 2000 international dollars[b]	Pardey *et al.* (2006), ASTI
Total trade	Agricultural export and import	Per cent of agricultural value added	FAO, WDI
Trade barrier	Average applied ad-valorem and ad-valorem equivalent agricultural protection	Tariff rate	CEPII, MAcMaps
Water	Water use by commodity	Million cubic metres	FAO, FEMISE

[a]The governance scores lie between −2.5 and 2.5, with higher scores corresponding to better quality of governance. For more details, see Kaufmann *et al.* (2007).
[b]The international dollar is a hypothetical unit of currency that has the same purchasing power that the US$ has in the US at a given point in time. The year 1990 or 2000 is often used as a benchmark year for comparisons that run through time.

Table A2. Descriptive statistics

	Mean	Standard deviation	Min	Max
Agricultural land	44.700	22.104	2.704	75.102
Agricultural machinery	5.230	4.706	0.450	21.104
Average holdings	3.060	3.481	0.251	20.220
Control of corruption	0.365	0.729	−0.881	1.690
Equipment import	5.571	4.422	0.440	19.780
Fertilizers consumption	1541.700	1131.009	50.504	4593.901
Fertilizers	4.212	9.751	0.001	62.121
Government effectiveness	0.434	0.816	−1.280	1.950
Human capital	6.110	1.781	3.010	9.403
Income	7508.800	6956.701	873.903	23641.302
Income Gini	37.710	6.713	23.021	55.041
Labour	28.100	49.940	0.052	289.703
Land	0.859	1.991	0.001	13.581
Land fragmentation	71.304	18.302	15.026	98.204
Land Gini	67.331	9.205	54.096	86.013
Land quality	27.025	22.714	6.012	100.017
Machines	31.863	69.543	0.016	434.530
Output	3.952	8.281	0.002	58.820
Political stability	−0.226	0.908	−2.492	1.280
Rain	157.012	157.923	7.034	478.127
R&D	316.303	723.234	8.702	3100.232
Total trade	5.081	12.206	0.122	132.604
Water	1615.912	5317.313	0.450	46146.093

Note: Summary statistics are computed over the period, countries, and commodities included in the sample. Definitions and units of measurements are provided in Table A1.

Corresponding author: Nadia Belhaj Hassine, Economic Research Forum, 7, Boulos Hanna St., Dokki 12311, Cairo, Egypt. E-mail: nbelhaj@erf.org.eg