European Market for Mercosur Agricultural Exports: An econometric study of commodity trade flows

J. Niemi;

Natural Resources Institute Finland (Luke), , Finland

Corresponding author email: jyrki.niemi@luke.fi

Abstract:

The European Union (EU), which represents one of the world’s largest markets for raw materials and agricultural products, with imports of more than €113 billion in 2015, is a particularly attractive and very sought-after market for Mercosur exporters. This paper provides new evidence on income and price elasticities of demand for agricultural exports from Mercosur countries to the EU. The econometric analysis of the study is conducted with a sample of annual data which covers Mercosur’s major commodity exports to the EU from 1988 to 2015. The results indicate that for the eight commodities analysed there is a relatively weak demand response to income and price changes in the EU. The policy implication is that trade policy measures in the form of tariff and non-tariff barriers are not very significant in changing the quantity of imports demanded. However, the results suggest that relative-price variations affect significantly the demand for Mercosur agricultural exports, implying that the exporter’s market share is influenced by price competitiveness.

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Keywords: agricultural trade, European Union, Mercosur, econometric models, cointegration

1 Introduction

The EU and Mercosur, a trade bloc in South America, agreed to continue negotiations on a long-stalled free trade accord by making fresh negotiating offers in May 2016. Mercosur encompasses Argentina, Brazil, Paraguay, Uruguay, as well as Venezuela, which is not currently participating in the negotiations. This was the first exchange of offers on goods and services between the two blocs since 2004, whereby the free trade negotiations originally began in 1999. The objective is to negotiate a comprehensive trade agreement, cutting customs duties, removing barriers to trade in services and improving rules related to public tenders, customs procedures, technical barriers to trade and protection of intellectual property. Stumbling blocks in the negotiations have been European demands to exclude agricultural exports — one of the most desired sectors for the South American countries — while Mercosur has been reluctant to lower its sizeable industrial tariffs, as requested by the EU.

The European Union (EU), which represents one of the world’s largest markets for raw materials and agricultural products, with imports of more than €113 billion in 2015, is a particularly attractive and very sought-after market for Mercosur exporters. The EU accounts for 35% of all Mercosur agricultural exports to the world. During the period between 1990 to 2015 Mercosur agricultural exports to the EU rose from €7 billion to €20 billion, showing an average annual growth rate of 4.2%. Over the years, Mercosur countries have also managed to increase their market share in the EU quite substantially. On the other hand, EU exports of agricultural products to Mercosur are not significant. Thus, the agricultural trade balance has clearly tilted in favour of Mercosur, with a trade surplus of over €18 billion in 2015.

Despite the success in penetrating to the EU market, the Mercosur countries has been concerned with the agricultural protection policy of the EU. The major irritant in EU-Mercosur agricultural trade relations has been tariffs and other discriminatory measures against Mercosur products, such as sugar and beef. Therefore, Mercosur countries have taken a special interest in encouraging the EU to liberalise its trade in agriculture. Mercosur countries hope that with trade liberalisation, the member countries will be able to improve their market access for agricultural products in the EU.
The objective of this study is to build a set of dynamic, theory-based econometric models which are able to capture both short-run and long-run effects of income and price changes on Mercosur’s major agricultural commodity exports to the EU, and which can be used for prediction and policy simulation under alternative assumed conditions. Many earlier studies (Bulmer-Thomas 2000, Flóres and Watanuki 2008, Burrell et al. 2011, Boulanger et al. 2016) have examined the effects of a free trade area formed by Mercosur and the EU on agriculture with aggregated commodities, but this paper attempts to increase our knowledge of the dynamic behavioural relationships underlying specific agricultural trade flows between Mercosur and the EU: beef, poultry, wheat, tomatoes, orange juice, sugar, coffee, and soybeans. The objective is to provide new estimates of short- and long-run elasticities of import and export demand for commodities exported from Mercosur countries to the EU.

The paper consists of five sections and it is organised as follows. Section two lays out the general theoretical and methodological framework employed by the study for modelling the dynamic relationships of commodity trade. Section three explains how the theoretical structures are implemented in dynamic econometric models in practice. Section four presents the empirical results of the estimated models constructed for eight agricultural commodities exported from Mercosur to the EU. In section five, we present a summary of our main conclusions.

2 Methodological framework

The estimation of import demand systems is derived from Armington’s (1969) model, where it is assumed that the same goods of different origins are imperfect substitutes within an importing country’s commodity market. Furthermore, in order to reduce to number of parameters to be estimated, the model assumes a constant elasticity of substitution (CES) for each product pair. Following the model, the importing decision is split into two stages. The solution to the utility maximisation problem for the first level of decision yields the overall demand schedules for commodity imports \( M \) of importer \( j \), given a commodity import price \( P \) and a level of constant dollar income \( Y \), and is expressed as

\[
M_j^d = k_1 Y_j \left( \frac{P_j}{D_j} \right)^{\varepsilon_{m}^d} \tag{1}
\]

where \( k_1 \) is a constant with expected sign \( k_1 > 0 \); \( D \) is the deflator; and \( \varepsilon_{m}^d \) is the price elasticity of import demand for goods \( M \). The income elasticity is equal to unity, a hypothesis that will later be tested.

Once the level of expenditures \( Y_j \) for the imported commodity \( M \) has been determined, the solution to the utility maximisation problem for how much of the commodity to purchase from alternative suppliers - let us say an exporter of interest \( i \) and its competitors \( m \), which refer each of the \( n-1 \) other foreign supplying countries, to market \( j \) whose corresponding export prices are \( P_{ij} \) and \( P_{mj} \) - may be expressed as

\[
X_{ij}^d = k_2 M_j \left( \frac{P_{ij}}{P_j} \right)^{\varepsilon_{x}^d} \tag{2}
\]

where \( X_{ij}^d \) is the quantity of the goods exported from country \( i \) to country \( j \), \( k_2 \) is a constant; \( P_{ij} \) is the price of the goods imported from country \( i \) to country \( j \); \( P_j \) is the average price of the goods imported to country \( j \); and \( \varepsilon_{x}^d \) is the relative-price elasticity of export demand.
Consider now the introduction of a tariff whose per-unit value is a specified amount into the import demand equation (1). The tariff raises the price of the product to \( (1+t)P \) in the geographic market \( j \). The resulting import demand schedule is

\[
M^d_j = k_j Y_j \left( \frac{(1+t)P_j}{D_j} \right)^{a_{ij}}
\]  

If export demand is proportional to the change in import demand in the geographic market and relative prices remain unaltered: that of the country of interest \( i \) and its competitor \( k \) is \( \frac{(1+t)P_i}{(1+t)P_k} \), the export demand schedules for the country \( i \) in the long-run dynamic equilibrium relationship implicit in equation (2) is

\[
X^d_{ij} = \tilde{a}_{ij} M_j \left( \frac{(1+t)P_j}{(1+t)P_i} \right)^{a_{ij}}
\]  

In other words, a change in the quantity demanded of the product as a result of tariff would cause a proportional change in the demand for the product supplied from foreign sources.

The empirical analysis of the study is based on econometric models which capture the dynamics underlying trade and price formation in commodity markets, and it is conducted by means of recently developed econometric concepts. Among these, the so-called ‘general to specific approach’ advocated by Hendry (1986) is applied in the context of data series whose (non-)stationary properties are investigated. Furthermore, the notion of cointegration (Engle and Granger, 1987) of a set of variables is analysed. The approach follows closely the modelling strategy developed in a series of papers by Davidson et al. (1978), Hendry (1986), Lord (1991), Urbain (1992), Banerjee et al. (1998) and Niemi (2003).

Given that economic time series often exhibit non-stationary stochastic processes, the econometric specification is conducted in a framework that allows for non-stationary but potentially cointegrated variables. The approach adopted is to convert the dynamic model into error correction formulation, and it is shown that this formulation contains information on both the short-run and long-run properties of the model, with disequilibrium as a process of adjustment to the long-run model. Equations specified in this manner allow the relevant economic theory to enter the formulation of long-run equilibrium in levels while the short-run dynamics of the equation are determined by growth rates.

Since the validity of the error correction specification requires the existence of a long-run relationship or cointegration between the variables concerned, the econometric analysis begins with the tests for the existence of a cointegrating vector. The first step in the analysis of cointegration is to determine the time series properties (i.e., the order of integration) of each variable, whether they have a unit root or not. Tests for unit roots are performed using the augmented Dickey-Fuller univariate tests. Having established the order of integration of each variable, tests for cointegration are undertaken and the nature of any cointegrating vectors explored. A formal test of cointegration is carried out following the residual-based approach proposed by Engle-Granger (1987) as well as the sequential testing procedure put forward by Perron (1988).

### 3 Econometric analysis of Mercosur’s agricultural exports to the EU

The econometric analysis of the study is conducted with a sample of annual data which covers Mercosur’s major commodity exports to the EU from 1988 to 2015. The analysis uses 4- and 5-digit
product-level data based on the Standard International Trade Classification (SITC). For the purpose of this study, the product headings are defined as follows: beef (SITC 011), poultry (SITC 01231, 01232), wheat (SITC 041), tomatoes (SITC 05440), orange juice (SITC 05910), sugar (SITC 061), coffee (SITC 071), and soybeans (SITC 08131).

Volume and value data on trade flows over the period 1988-2015 are obtained from Eurostat. Volume data is compiled in metric tons, and value data in thousands of euros. The transaction value is the value at which goods were sold by the exporter, and includes the cost of transportation and insurance, and freight to the frontier of the importing country (c.i.f. valuation).

The unit prices of EU imports ($P_{EU}$), and unit prices of exports by an individual Mercosur country ($P_{j}$), are derived by dividing value by volume. These individual product unit values are subject to an error in measurement. If import declarations are inaccurate, customs data may be incorrect. Scobie and Johnson (1975) have shown that, if the observed value and the volume data contain errors of measurement for actual transactions, the estimated elasticity of substitution will be biased towards zero. Moreover, unit values suffer from the traditional f.o.b./c.i.f. valuation problems. Elasticity estimates are based on c.i.f. prices, which, because they include changes in trade resulting from transportation and distribution costs or from tariffs, do take into account all price differences between suppliers to the ultimate consumer (Lord, 1991). Therefore, in this ‘standard’ formulation the observed real prices in exporting country $i$ assume fixed transfer costs.

The gross domestic product (GDP) index and the consumer price index (CPI) are used as a measure of economic activity ($Y_{EU}$) and price deflator ($D_{EU}$) of the EU, respectively. The source of the data is the International Financial statistics database of the International Monetary Fund (IMF).

Models such as equation (1) are usually specified in log-linear form by assuming that standard trade theory relates exports and imports to explanatory variables through a multiplicative form which can be derived within a cost minimisation framework (Urbain, 1992). The first-order stochastic difference equation as a logarithmic function of the theoretical relationship in (1) is, therefore, expressed as

$$\ln M_{j,t} = \alpha_{0} + \alpha_{1} \ln Y_{j,t} + \alpha_{2} \ln Y_{j,t-1} + \alpha_{3} \ln \left( \frac{P_{j}}{D_{j}} \right)_{t} + \alpha_{4} \ln \left( \frac{P_{j}}{D_{j}} \right)_{t-1} + \alpha_{5} \ln M_{j,t-1} + v_{it} \quad (5)$$

where the expected signs are $\alpha_{1}, \alpha_{2} > 0; \; \alpha_{3}, \alpha_{4} < 0; \; \text{and} \; 0 < \alpha_{5} < 1$. The lags in the model are specified as the maximum to be expected in the light of the nature of import demand and the evidence of previous econometric studies. The maximum lag length for annual time-series data is usually equal to one on the hypothetical basis that economic agents are characterised by one-year planning horizons.

The results of the cointegrating regressions in Appendix A show that demand for commodity imports in the EU market (ln $M_{EU}$) has a steady-state response to the domestic economic activity ($\ln Y_{EU}$), and a transient response to the constant dollar price of imports ($P/D$).

Transformation of equation (3) to incorporate an ECM driven by ln $Y_{EU}$ and (ln $P/D_{j}$) to the equation (3), with an additional lagged variable of ln $Y_{EU}$, results in the following import demand specification:

$$\Delta \ln M_{j,t} = \alpha_{0} + \alpha_{1} \Delta \ln Y_{j,t} + \delta_{1} \ln Y_{j,t-1} + \alpha_{3} \Delta \ln \left( \frac{P_{j}}{D_{j}} \right)_{t} + \delta_{3} \Delta \left( \frac{P_{j}}{D_{j}} \right)_{t-1} + \delta_{5} \ln \left( \frac{M_{j}}{Y_{j}} \right)_{t-1} + v_{it} \quad (6)$$
where $\delta_2 = (\alpha_1 + \alpha_2 + \alpha_3 - 1)$, $\delta_3 = (\alpha_3 + \alpha_4)$, and $\delta_5 = (\alpha_5 - 1)$. The expected signs of the coefficients are $\alpha_1 > 0$, $\delta_2 > \delta_3 > -1 < \delta_5 < 0$, and $\alpha_3, \delta_4 < 0$. The fifth term of the equation, $\ln (M/Y)_{t-1}$ is called the error correction term, while $\delta_5$ is the feedback coefficient. The error correction term captures the adjustment toward the long-run equilibrium. If $\delta_5$ is statistically significant, it states what proportion of the disequilibrium in $\Delta \ln M_j$ in one period is corrected in the next period.

The long-run dynamic solution of a single-equation system generates a steady-state response in which growth occurs at a constant rate, say $g$, and all transient responses have disappeared (Currie, 1981, Lord, 1991). With growth rates of domestic economic activity and import demand, $\Delta \ln Y_j = g_1$ and $\Delta \ln M_{jj} = g_2$, respectively, the long-run dynamic equilibrium solution of equation (3), in terms of the original (anti-logarithmic) values of the variable, is

$$M_j = k_1 Y_{jt}^{\delta_1/d_5} \left( P_j / D_j \right)^{\delta_1/d_5}$$

(7)

where $k_1 = \exp \{-\alpha_0 + (1-\alpha_1)g_1/\delta_5\}$. Equation (5) encompasses the static equilibrium solution when $g_1 = 0$. The income elasticity of import demand is expressed as $\epsilon_Y^{\delta_1} = 1 - (\delta_2/\delta_5)$. The price elasticity of import demand is $\epsilon_{Pj}^{\delta_5} = -\delta_0/\delta_5$.

In terms of the general stochastic difference specification, the export demand relationship in (2) is expressed as

$$\ln X_{ijt}^d = \beta_0 + \beta_1 \ln M_{jt} + \beta_2 \ln M_{j,t-1} + \beta_3 \ln \left( \frac{P_{ij}}{P_j} \right) + \beta_4 \ln \left( \frac{P_{ij}}{P_j} \right)_{t-1} + \beta_5 \ln X_{ij,t-1}^d + v_{2t}$$

(8)

where the expected signs of the coefficients are $\beta_1, \beta_2 > 0$; $\beta_3, \beta_4 < 0$; and $0 < \beta_5 < 1$. The dynamics for the export demand relationship is assumed to be of relatively small order, and can therefore be restricted to cases where the lagged values of the variables are of one year. The Lagrange multiplier (LM) tests are again performed for omitted higher lagged variables.

The results of the cointegrating regressions suggest that in most cases the demand for exports of a Mercosur country $i$ ($\ln X_{ijt}^d$) has a steady-state response to the import demand of the EU ($\ln M_{EUt}^d$), and a transient response to the relative price of the EU market ($\ln P_{ij}/\ln P_{EU}$). In the second case, demand for exports from country $i$ to country $j$ ($X_{ij}$) has a steady-state response both to the import demand ($M_j$), and to the relative price ($P_{ij}/P_j$) of that market. In other words, $X_{ij}, M_j, P_{ij}/P_j$ are cointegrated. An ECM is obtained from the cointegration regression of $X_{ij}$ on $M_j$ and $P_{ij}/P_j$. The following transformation of (6) incorporates an ECM driven by import demand $M_j$:

$$\Delta \ln X_{ij}^d = \beta_0 + \beta_1 \Delta \ln M_{jt} + \gamma_2 \Delta \ln \left( \frac{P_{ij}}{P_j} \right) + \gamma_3 \ln \left( \frac{P_{ij}}{P_j} \right)_{t-1} + \gamma_4 \ln \left( X_{ij}^d / M_{j,t-1} \right) + v_{2t}$$

(9)

where $\gamma_2 = \beta_3$, $\gamma_3 = \gamma_2 = (\beta_3 + \beta_4)$, and $\gamma_4 = (\beta_4 - 1)$. The expected signs of the coefficients are $\beta_1, \gamma_2 > 0$, $\gamma_3 < 0$, and $-1 < \gamma_4 < 0$. The relative price term in the foregoing specification have been so transformed as to nest the ‘differences’ formulations of the variable in the levels form of the equation. The error correction term, $\gamma_4 \ln (X_{ij}/M_{ij,t-1})$, measures divergences from the long-run equilibrium and corrects for previous non-proportional responses in the long-run dynamic growth of export demand. Since in dynamic equilibrium $\Delta \ln M_{jj} = g_2, \Delta \ln X = g_3$ and $\Delta \ln (P_{ij}/P_j) = 0$, it follows that the solution of (7), in terms of the original values of the variable, is

$$X_{ij}^d = k_2 M \left( \frac{P_{ij}}{P_j} \right)^{-\gamma_3/\gamma_4}$$

(10)
where \( k_2 = \exp \{[-\beta_0 + (1-\beta_1)g_2]/\gamma_4 \} \). Therefore, export demand is assumed to have a unitary elasticity with respect to the level of import demand in the geographic market. The price elasticity of export demand is expressed as \( \varepsilon_x = -\gamma_3/\gamma_4 \).

Where the long-run response between the export demand of a country \( i \) and imports of its trading partner \( j \) is not necessarily proportional, an additional term (explanatory variable for imports of a country \( j \) lagged by one period) is introduced into the equation (6).

The foundations of the error correction model (ECM) specification, used in equations (4) and (8), rest on the seminal work of Sargan (1964). The ECM specification can be derived as a simple reparameterisation of a general autoregressive distributed lag (ADL) model. The idea of incorporating the dynamic adjustment to steady-state targets in the form of error-correction terms, suggested by Sargan and developed by Davidson et al. (1978), among others, offers the possibility of revealing information about both short-run and long-run relationships.

4 Estimation results

4.1. The import demand functions

The short- and long-run responsiveness of agricultural commodity imports to changes in incomes and own-prices in the EU are summarised in Table 1. Coefficient signs and magnitudes are acceptable in terms of a priori expectations. The models also track the sizes and the directions of changes in the volume of EU agricultural imports fairly well. Considering that the equation explains the rate of changes in the import volumes, the \( R^2 \) values ranging from 0.41 to 0.69 can be considered quite satisfactory.

The results for own-prices elasticities indicate that they are statistically different from zero in six out of the eight commodities, and, of these, one is significant at the 1% level, one at the 5% level, and four at the 10% level. The estimated income elasticities have expected positive signs and are significantly different from zero at the 10% level in the equation for all commodities, excluding sugar. All the coefficients of the lagged error-correction terms appear highly significant at 1-percent level. Therefore, the deviation from the equilibrium level of the import demand due to random shocks represents a significant determinant of its short-run dynamic behaviour.

The coefficient estimates on the own-price terms confirm the expectation that demand for commodity imports in the EU is relatively inelastic with respect to price. The price elasticities range from -0.08 to -0.34 in the short-run, and from -0.11 to -0.72 in the long run. The policy implication of this fact is that exchange rate policies and commercial policy intervention measures in the form of tariff and non-tariff barriers to trade would not be very effective in changing the quantity of imports demanded. Soybeans has the largest long-run price elasticity (\( \varepsilon = -0.72 \)) and the remaining five have elasticities less than 0.5 in absolute terms. Coffee has the lowest long-run price elasticity (\( \varepsilon = -0.11 \)).

Beef and sugar did not show the expected sign of price elasticity, though neither was statistically significant. Beef and especially sugar has been highly protected agricultural products in the EU internal markets. The standard tariffs and additional import duties for sugar have been so prohibitive that almost all sugar imports are from developing countries that receive preferential treatment from the EU. Thus, price and income elasticity estimation for sugar may be an impossible task.

The long-run income elasticities are in a range between 0.09 for wheat and 1.02 for soybeans. Orange juice is found to have a unitary elasticity with respect to income. Beef has income elasticity
close to unity, and wheat and tomatoes have elasticities significantly less than unity. The large differences in the income elasticity have implications for sales by exporters. Soybeans imports have been sensitive to income changes. Thus, soybeans exports have a considerably stronger growth potential in the EU than other commodities, because of a strong response of buyers in the EU to improvements in their real income.

The high income elasticity of soybeans is supported by a strong demand for its use as an ingredient in compound animal feed. The level of demand for compound feeds depends on the livestock industry, and the livestock industry depends in turn on the level of demand for meat and other livestock products.

Overall, the results suggest that agricultural commodity imports of the EU are not very sensitive to income changes and are considered necessary goods in the sense that demand increases slower than economic activity goes up. This means a relatively weak growth potential for the selected commodities in the EU market. At the same token, imports of these commodities are not susceptible to larger swings of demand during business cycles, either. The trade-weighted average long-run income elasticity of import demand across commodities is relatively low (≈0.53).

The findings are consistent with the earlier studies. The low income and price elasticity of demand for primary commodities have been recorded in many studies covering a wide range of commodities.

4.2. The export demand functions

The elasticity estimates of export demand equations for the major commodity exports of Mercosur to the EU are reported in Table 2. The signs and magnitudes of the estimated coefficients are broadly in line with theoretical expectations. Furthermore, the models explain the changes in the volume of Mercosur agricultural exports to the EU rather accurately. Goodness of fits is acceptable with an $R^2$ in a range between 0.35 and 0.76. The models also pick up quite well the turning points and rapid rises in export demand.

As expected, relative price movements affect significantly the trade flows of all commodities, implying that exporter’s market share has been influenced by price competitiveness. Relative prices are statistically different from zero in 9 out of the 14 trade flows, and, of these, two are significant at the 1% level, three at the 5% level, and four at the 10% level.

The sizes of relative price coefficients, of course, differ by commodity as well as by source of supply in each commodity. The short-run relative price elasticity of export demand range from −0.34 to -0.67, and the long-run elasticity from -0.25 to -2.25. In other words, there is a great deal of variation in the export performance between different commodities and among individual Mercosur countries. Therefore, care should be exercised in generalisations about the price elasticities of demand for the region’s commodity exports.

The observed differences in relative-price coefficients by trade flow reflect the dynamic aspect of the EU agricultural trade, in which particular trade flows rise and fall with price competition. Among the trade flows under examination, the export demand for Argentinian soybeans is the least sensitive to relative price changes, followed by Brazilian coffee exports. In contrast, the relative-price coefficients of Brazilian orange juice and Argentinian exports are exceptionally large, -2.25 and -2.09, respectively.

These findings, combined with the result of import price elasticity in Table 2, indicate that, although agricultural imports are relatively insensitive to price changes on a commodity basis, once the total amount to be spent for imports of a commodity is determined then the EU importers seek cheaper
products, so that price competition among suppliers is inevitable. On the other hand, the sharp contrast of relative price coefficients in the same commodity justifies the assumption that importers distinguish agricultural products by place of production, even though the products are called by a common commodity name.

The estimation results also confirm the assumption that export demand for commodities from Mercosur has, in general, more or less proportional response to changes in the level of EU import. Therefore, at given relative-price levels, any increase or decrease in commodity imports by the EU would be reflected in an almost equivalent percentage change in its demand for exports from Mercosur countries. In other words, the market share of the country does not change unless relative prices change in homothetic demand.

However, if the estimated coefficient of the import response variable is significantly greater than unity, it is a good indication for an exporting country that its exports can expand more than others and its share increase as EU market grows. Among the selected commodity trade flows, orange juice and beef from Brazil have clearly more than proportional response to changes in the level of EU imports.

5 Conclusions

This paper has attempted to apply a reasonably flexible data determined, dynamic model to estimate the short-run and long-run effects of changes in income and prices on agricultural commodity trade between Mercosur and the EU. Therefore, a modelling approach based on the error correction mechanism (ECM) was used in order to emphasise the importance of dynamics of trade functions. Application of a series of diagnostic tests supported the use of this approach. Econometric models were constructed for eight agricultural commodities – beef, poultry, wheat, tomatoes, orange juice, sugar, coffee, and soybeans - exported from Mercosur to the EU.

The results for the estimated import demand functions suggest that there is a relatively weak demand response to income changes in the EU. The results also demonstrate the inelastic nature of price responses in the EU demand for the imported commodities. The policy implication of this fact is that trade policy measures in the form of tariff and non-tariff barriers are not very significant in changing the quantity of imports demanded.

In the case of beef and sugar, imports would likely increase far more than those of any other commodity if import protection is abolished. However, due to the high level of protection, which distorts import demand responses, we were not able to get statistically significant price elasticity estimates for beef and sugar.

The coefficient estimates of the export demand functions indicate that relative-price variations affect significantly the demand for Mercosur commodity exports by the EU, implying that exporter’s market share is influenced by price competitiveness. Furthermore, the sharp contrast of relative price coefficients in the same commodity across countries justifies the assumption that importers distinguish agricultural products by place of production, even though the products are called by a common commodity name.

The estimated equations of the individual commodities of this study can be used to simulate the effects of changes in economic environment and trade policies on agricultural commodity trade between Mercosur and the EU. More specifically, the set of dynamic econometric models
developed in the preceding chapter can be used to examine the agricultural commodity trade effects of a free trade area formed by the EU and Mercosur. Of interest are both the short-run effects and the long-run effects after the full adjustment has taken place. Furthermore, the estimated equations can be utilized to quantify how changes in the exchange rate affect the volume and prices of Mercosur agricultural exports to the EU.

6 References


7 Tables

Table 1. Short-run and long-run elasticities of import demand in the EU for selected commodities.

<table>
<thead>
<tr>
<th>Commodity</th>
<th>Price elasticity</th>
<th>Income elasticity</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Short-run</td>
<td>Long-run</td>
</tr>
<tr>
<td>Beef</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Poultry</td>
<td>-0.22</td>
<td>-0.37</td>
</tr>
<tr>
<td>Wheat</td>
<td>-0.16</td>
<td>-0.32</td>
</tr>
<tr>
<td>Tomatoes</td>
<td></td>
<td>-0.17</td>
</tr>
<tr>
<td>Orange juice</td>
<td>-</td>
<td>-0.35</td>
</tr>
<tr>
<td>Sugar</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Coffee</td>
<td>-0.08</td>
<td>-0.11</td>
</tr>
<tr>
<td>Soybeans</td>
<td>-0.34</td>
<td>-0.72</td>
</tr>
</tbody>
</table>

Note: - Not significant at the 10% level

Table 2. The short-run and long-run responsiveness of EU’s agricultural imports from the Mercosur countries to changes in relative prices and in the level of EU imports.

<table>
<thead>
<tr>
<th>Commodity</th>
<th>Exporter</th>
<th>Relative price elasticity of export demand</th>
<th>Response to changes in the level of EU imports</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>Short-run</td>
<td>Long-run</td>
</tr>
<tr>
<td>Beef</td>
<td>Argentina</td>
<td>-</td>
<td>-2.09</td>
</tr>
<tr>
<td></td>
<td>Brazil</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>Uruguay</td>
<td>-0.34</td>
<td>-0.58</td>
</tr>
<tr>
<td>Poultry</td>
<td>Argentina</td>
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<td>-</td>
</tr>
<tr>
<td></td>
<td>Brazil</td>
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<td>-0.75</td>
</tr>
<tr>
<td>Wheat</td>
<td>Argentina</td>
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Note: - Not significant at the 10% level