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# Brazilian Import Demand of Dairy Products with Emphasis in the Mercosul Context

*Demanda brasileira de importações de laticínios com ênfase no contexo do Mercosul*

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## Abstract

Brazil is the largest milk producer in South America, with an increasing trend in the last 20 years which allowed the country to start exporting some output. However, it has not become self-reliant and continues to be a major importer in the region, the most relevant partner for Argentina and Uruguay in Mercosul. In this paper, the Brazilian import demand of dairy products is estimated by using a source differentiated demand system framework that is derived from consumer theory. This model takes into account the countries of origin (i.e., Argentina, Uruguay and other grouped competitors) and various dairy aggregated items that compete in the Brazilian market. The expenditure elasticities, and the own and cross-price elasticities are computed with the estimated parameters of the demand system. Interpretation of results follows as well as some conclusions, including the need of further studies and usefulness of this type of research.

## Keywords

Almost Ideal Demand System, Source Differentiated, Price-Elasticity, Expenditure-Elasticity, Argentina, Uruguay.

**JEL Codes** D12, F14, L66.

## Resumo

*O Brasil é o maior produtor de lácteos da América Latina e, nos últimos 20 anos, deixou de ser um importador líquido de produtos lácteos para ser um exportador. No entanto, não se tornou autossuficiente e o Brasil continua a ser um grande importador na região, e é o parceiro mais relevante para a Argentina e o Uruguai no Mercosul. Neste trabalho, a demanda de importação brasileira de produtos lácteos é estimada usando uma estrutura de sistema de demanda diferenciada de origem que é derivada da teoria do consumidor. Esse modelo leva em conta os países de origem (ou seja, Argentina, Uruguai e outros concorrentes agrupados) e vários itens agregados de lácteos que competem no mercado brasileiro. elasticidades preço próprias e cruzadas, além da elasticidade de demanda em relação ao gasto total são calculadas com os parâmetros estimados do sistema de demanda. As conclusões sobre o potencial competitivo dos exportadores envolvidos neste mercado são obtidas, com especial ênfase para os parceiros do Mercosul.*

## Palavras-chave

*Sistema de demanda quase ideal, Fonte Diferenciada, Elasticidade de preço, Elasticidade das despesas, Argentina, Uruguai.*

**Códigos JEL** *D12, F14, L66.*

# 1 **Introduction**

South American countries constitute a significant proportion of the international demand for regional dairy exports, with Brazil and Venezuela leading the list. Despite the fact that Brazil is the largest milk producer in the region with more than 33 million tons per year,[1] volumes are still not sufficient to satisfy all domestic needs. The very favourable evolution of its milk production in the last decades allowed the country to venture into exporting some local output (Carvalho, 2010; Depetris Guiguet *et al*., 2010) and more recently began discussing a Brazilian milk export policy. While this meant a greater capacity to satisfy domestic demand, and in turn, the possibility to compete in certain international markets, Brazil's imports remain significant.

In fact, initially, the increase in dairy production was paralleled by a continuous drop in its dairy imports, particularly in the period 1995-2006. However, Brazilian imports started to grow again since 2007, as its exports did. Thus, at the same time that the country succeeded in exporting higher volumes, its dairy imports were reactivated to satisfy its local demand. In this way, Brazil's dairy imports reached in 2015 similar levels to those of 2000-2001.

The more significant deficit in consumption occurs with whole milk powder (WMP), skim milk powder (SMP) and cheese, the three categories that conform the bulk of dairy imports, as fluid milk imports are insignificant. As for their relative weights, even though they present a pronounced variation every year, WMP imports more than double each of the other two. The latest figures for 2017 indicate imports of 73,000 tons of WMP, 31,000 tons of SMP and 32,000 tons of cheese.[2]

As a full member of the Mercosul agreement, vicinity as well as the preferential tax treatment facilitate Brazilian dairy imports from Argentina and Uruguay. In WMP imports their participation was 37% and 58% respectively for 2017; in SMP 45% and 34%; and in cheese imports 49% and 37%.[3]

Therefore, Brazilian dairy demand continues to be one of the main channels where Argentina and Uruguay manage to sell a majority of their

1 https://www.clal.it/en/?section=stat_brasile.

2 https://www.clal.it/en/?section=stat_brasile.

3 https://www.clal.it/en/?section=stat_brasile.

dairy exports. Notwithstanding the commercial advantages given by distance and their Mercosul membership, both countries must compete: between themselves, with other dairy exporters and the growing domestic production in Brazil. Under this scenario, the aim of the present paper is to understand the elasticities of the demand of imports and its relation with the competitive position of exporting countries, focusing on Mercosul exporters. Despite the importance of the Brazilian market for Mercosul dairy exports, the related economic literature is very scarce. One of the closest studies is the recent contribution of Pinha *et al*. (2016), who use a residual demand approach to estimate the market power of Argentina and Uruguay in the Brazilian import market for milk powder. However, this paper does not aim to study the Brazilian demand behavior but it instead aims to determine whether there are indications of market power in the exporter's pricing. In the present research, we focus on the study of the behavior of Brazil as an importer of dairy products and its implications for the competitive positioning of the exporting countries.

In general terms, the micro-based models of import demand follow Armington's sources differentiation approach (Armington, 1969). Beyond the fact that Armington's traditional model and its extensions (under CES preferences) can yield a parsimonious demand system for its econometric treatment, the recent empirical literature has turned to more flexible approaches but are still consistent with formal economic theory. In particular, two approaches are used to estimate import demand systems that have had significant diffusion in the last 15 years: the first is developed from consumer theory and the second uses production theory.

The consumer demand theory treats imports as products that fall directly into the consumers' utility function, whereas production theory treats imports as inputs purchased to be transformed by companies in the importing country. In the first case, the demand for imports derives from the process of maximizing the utility of consumers in the purchasing country. In the second case, the demands are derived from the process of profit maximization or cost minimization of the companies that use the imports as an input of the production process of another, more elaborate good. Consequently, the empirical applications are made on imported goods that are clearly used as inputs by the importing country (Washington and Kilmer, 2002, Muhammad, Jones and Hann, 2004, Feleke and Liu, 2005, Muhammad, 2008, McLaren and Zhao, 2009).

Given that processed products are traded by international firms from each country (Depetris Guiguet *et al*., 2010), in this study we follow the consumer theory approach.

Deaton and Muellbauer's (1980) Almost Ideal Demand System (AIDS) and its different variants have gained great importance in applied research to estimate both domestic demand and imports, mainly because of its flexibility and consistency with economic theory (Barnett and Serletis, 2008). The AIDS specification that is applied to import demand in its standard form (*e.g*. Balagtas *et al*., 2006 and Susanto *et al*., 2008) does not provide information on competitiveness among exporting nations. Therefore, Yang and Koo (1994) propose a version with differentiation by origins, which they use to estimate Japanese meat import demand. A number of papers have also adopted this specification and applied it to different international agri-food markets, such as: Andayani and Tilley (1997), Fabiosa and Ukhova (2000), Lee *et al*. (2008). In particular, for dairy products Ramirez and Wolf (2008) studied Mexican demand for imports, considering the categories of milk powder, cheese and other dairy products, and three sources: the United States, the European Union and Others (aggregates).

Based on this approach of (almost ideal) demand systems differentiated by source countries, this paper studies the Brazilian demand for dairy imports, where Argentina and Uruguay maintain a majority position. Three dairy products are considered: Whole Milk Powder (WMP), Skim Milk Powder (SMP) and Cheese (all forms of cheese are aggregated in one category). As source countries, we take Argentina, Uruguay and other international competitors (Others) as an aggregate origin because we are able to derive the cross effects to detect the degree of competition.

## 2 **Theoretical Model**

To model import demand, we assume separability in the Brazilian preferences between domestic and imported dairy products. Additionally, we adopt the Armington hypothesis of import differentiation by country of origin (Armington, 1969). Under these theoretical assumptions, this paper uses a demand specification that is derived from consumer theory. Specifically, we propose a demand system approach for aggregate dairy products, differentiating them by the origins of imports (*i.e.* the exporting countries).

The Source Differentiated Almost Ideal Demand System (SDAIDS) was proposed formally by Yang and Koo (1994) to analyze the Japanese meat import demand. This was based on Deaton and Muellbauer's (1980) contribution about the specification of flexible demand system, which is consistent with the microeconomics consumer theory.

The demand system for an importing country is derived by assuming that it has Price Independent Generalized Logarithmic (PIGLOG) preferences. This implies that the expenditure function is given by

$$\ln E_m(\mathbf{p},U_m) = (1-U_m)\ln a(\mathbf{p}) + U_m \ln b(\mathbf{p}) \tag{1}$$

where $\mathbf{p}$ represents the $N_m$ -dimensional vector of import prices of different sources, $U_m$ is a utility index, and the functions $a(\mathbf{p})$ and $b(\mathbf{p})$ are of the form

$$\ln a(\mathbf{p}) = \alpha_0 + \sum_i \sum_h \alpha_h^i \ln p_h^i + \frac{1}{2}\sum_i \sum_j \sum_h \sum_k \lambda_{hk}^{ij} \ln p_h^i \ln \ln p_k^j \tag{2}$$

and

$$\ln b(\mathbf{p}) = \ln a(\mathbf{p}) + \beta_0 \prod_i \prod_h \left(p_h^i\right)^{\beta_h^i} \tag{3}$$

where $\alpha_0$, $\alpha_h^i$, $\lambda_{nk}^{ij}$, $\beta_0$ and $\beta_h^i$ are parameters with $i,j = 1,\ldots,N$, $h = 1,\ldots,H_i$ and $k = 1,\ldots,K_j$. The superscripts $i$ and $j$ denote the imported good (*i.e.* a specific dairy product) and the subscripts $h$ and $k$ indicate the source of imports. This notation formalizes the hypothesis that two different products ($i \neq j$) may have different countries of origin; that is, $H_i$ sources for good $i$ and $K_j$ for the good $j$. Then, replacing (2) and (3) in (1), the expenditure function takes the following expression

$$\ln E_m(\mathbf{p},U_m) = \alpha_0 + \sum_i \sum_h \alpha_h^i \ln p_h^i + \frac{1}{2}\sum_i \sum_j \sum_h \sum_k \lambda_{hk}^{ij} \ln p_h^i \ln p_k^j + \\ + \beta_0 U_m \prod_i \prod_h \left(p_h^i\right)^{\beta_h^i} \tag{4}$$

Applying *Shephard's* lemma, and differentiating the function (4) respect to the logarithm to the price of good $i$ imported from country $h$ (*i.e.* In $p_n^i$), we obtain the budget shares equations ($w_n^i$) as a function of the prices vector $\mathbf{p}$ and the utility level $U_m$. These equations are given by

$$w_h^i = \alpha_h^i + \sum_j \sum_k \gamma_{hk}^{ij} \ln p_k^j + \beta_h^i \beta_0 U_m \prod_i \prod_h \left(p_h^i\right)^{\beta_h^i - 1} \tag{5}$$

where $w_h^i = p_h^i q_h^i / E_m$ and $\gamma_{hk}^{ij} = (\lambda_{hk}^{ij} + \lambda_{kh}^{ji}) / 2$. Taking into account that the total expenditure in imports is equal to the expenditure function valued at the utility maximum point ($E_m \equiv E$), we can remove the utility index $U_m$ from (5). Specifically, by resolving (4) for $U_m$, we have the indirect utility function $V_m(\mathbf{p},E)$, which can be used in (5) to take the place of $U_m$. This enables us to obtain a system of share equations, corresponding to the Marshallian imports demand; that is,

$$w_h^i = \alpha_h^i + \sum_{j=1}^{N} \sum_{k=1}^{K_j} \gamma_{hk}^{ij} \ln p_k^j + \beta_h^i \ln\left(\frac{E}{P}\right) \tag{6}$$

with $i = 1,\ldots,N$, $h = 1,\ldots,H_i$ and

$$\ln P = \alpha_0 + \sum_i \sum_h \alpha_h^i \ln p_h^i + \frac{1}{2} \sum_i \sum_j \sum_h \sum_k \lambda_{hk}^{ij} \ln p_h^i \ln p_k^j \tag{7}$$

The properties of linear homogeneity of the expenditure function, the symmetry derived from second order conditions, and the adding-up among share equations, imply that a set of parameters constraints for the system (7) are required to have consistency with the basic microeconomic postulates. More precisely, these constraints are:

Adding-up:

$$\sum_i \sum_h \alpha_h^i = 1, \ \sum_i \sum_h \gamma_{hk}^{ij} = 0, \text{ and } \sum_i \sum_h \beta_h^i = 0 \tag{8}$$

Homogeneity:

$$\sum_j \sum_k \gamma_{hk}^{ij} = 0 \tag{9}$$

Symmetry:

$$\gamma_{hk}^{ij} = \gamma_{kh}^{ji} \tag{10}$$

Another constraint that must be satisfied is related to the curvature of the utility function, and is the so-called negativity condition (Barten and Geyskens, 1975). Specifically, the required quasi-concavity for the utility function implies that the Slutsky matrix (**S**) must be negative semi-definite and, therefore, the compensated Hicksian demands will be non-increasing with respect to their own prices (Holt and Goodwin, 2009). Following to Holt and Goodwin (2009), and Feleke and Liu (2005) in their more empirical version, for the case of import demand system, the negativity condition must be satisfied if the following matrix is negative semi-definite:

$$[\mathbf{S}]_{hk}^{ij} = \gamma_{hk}^{ij} + \beta_h^i \beta_k^j \ln(E/P) - \delta_{hk}^{ij} w_h^i + w_h^i w_k^j \tag{11}$$

where $[\mathbf{S}]_{hk}^{ij}$ is the element of *ih*-row and *jk*-column of matrix **S**, and $\delta_{hk}^{ij}$ is the delta Kronecker between the good *i* from source *h* and the good *j* from source *k*.

For a certain imported good *i* to be exported by country *h*, the expenditure elasticity (*i.e.* the percentage change in the quantity demanded of the good *i* with origin *h* as a result of a change of 1 percent in the total expenditure in dairy imports) is obtained from (6). In particular, if we note that $\partial w_h^i / \partial \ln E = \beta_h^i$ and, therefore, $\partial \ln w_h^i / \partial \ln E = \beta_h^i / w_h^i$, then, considering that the differential $d \ln w_h^i$ can be decomposed as the sum $d \ln p_h^i + d \ln q_h^i - d \ln E$, we have

$$\eta_{i,h}^E \equiv \frac{\partial \ln q_h^i}{\partial \ln E} = 1 + \frac{\beta_h^i}{w_h^i} \tag{12}$$

In the same way, the price elasticities can be derived from the share equations. Specifically, the uncompensated price elasticity (*i.e.* Marshallian) for all demand system is equal to minus delta Kronecker ($\delta_{hk}^{ij}$) plus the partial derivative of the logarithm of budget share with respect to the logarithm of price (Holt and Goodwin, 2009). For the sources differentiated demand system, the Marshallian price elasticity of good *i* from source *h* with respect to the price of good *j* from source *k* will be $-\delta_{hk}^{ij} + \partial \ln w_h^i / \partial \ln p_k^j$ where $\delta_{hk}^{ij} = 1$ if $i = j$ and $h = k$, and $\delta_{hk}^{ij} = 0$ if $i \neq j$ and/or $h \neq k$. Then, from (6) and (7), we have the following formula for uncompensated price elasticity:

$$\eta_{hk_U}^{ij} = \delta_{hk}^{ij} + \frac{\gamma_{hk}^{ij}}{w_h^i} - \frac{\beta_h^i}{w_h^i}\left( \alpha_k^j \sum_x \sum_y \lambda_{yk}^{xj} \ln p_y^x \right) \tag{13}$$

where $\gamma_{hk}^{ij}$ is the price coefficient, $\beta_h^i$ is the expenditure coefficient, and the expression in brackets arises from the derivative $\partial \ln P / \partial \ln p_k^j$ that is computed from(7). Using the share equation, this expression is equal to

$(w_k^j - \beta_k^j \ln(E / P))$, which is usually used in the empirical literature. Therefore, from (13) the Marshallian own-price and cross-price elasticities can be known. In this context, these elasticities are interpreted as the percentage change in the quantity demanded of the good *i* imported from *h*, due to a change of 1 percent in the price, while maintaining a constant nominal total outlay in dairy imports.

From the Slutsky equation, we can obtain the compensated (Hicksian) price elasticity, which can be written as:

$$\eta_{hk_C}^{ij} = \eta_{hk_U}^{ij} + w_k^j \eta_{i,h}^E \tag{14}$$

The Hicksian price elasticity $\eta_{hk_C}^{ij}$ measures the percentage change in the quantity demanded of *i* from *h* as a response to a change of 1 percent in the price of good *j* imported from *k*, holding the real total outlay in dairy imports.

Taking into account that the equations (6) and (7) specify a non-linear demand system, in this paper we follow the Deaton and Muellbauer (1980) approach, who proposed replacing the non-linear price index (7) with a linear approximation, obtaining a linear system of the budget share equations(*LA-AIDS*). In particular, Deaton and Muellbauer (1980) suggest the Stone price index, given by

$$\ln \tilde{P} = \sum_j \sum_k w_k^j \ln p_k^j \tag{15}$$

Using (15) we obtain a more parsimonious system for empirical application. This approach is commonly adopted in the literature of import demand systems estimation (Seale *et al.*, 2003; Gil *et al.*, 2004; Feleke and Liu, 2005; Ben Kaabia and Gil, 2007; Wan *et al.*, 2010; Nzaku *et al.*, 2012; Lee *et al.*, 2014, and others). Once a linear approximation is adopted, the price elasticities (13) and (14) become:

Uncompensated price elasticities (Marshallian):

$$\eta_{hk_U}^{ij} = \delta_{hk}^{ij} + \frac{\gamma_{hk}^{ij}}{w_h^i} - \beta_h^i \frac{w_k^j}{w_h^i} \tag{16}$$

Compensated price elasticities (Hicksian):

$$\eta_{hk_C}^{ij} = \delta_{hk}^{ij} + \gamma_{hk}^{ij} + w_k^j \tag{17}$$

# 3 **Methodology**

## 3.1 **Data and Variables**

For the empirical analysis, we take customs data from a private international database (Penta-Transaction, 2014) that collects detailed trade and is mainly focused on South-American countries. In general, for each trade flow (export or import) the nomenclature, the date, the country of origin and destination, the customs office of departure-arrival, the quantity traded, the F.O.B value for exports and the C.I.F value for imports are reported. Depending on the specific country (Customs), extra information is included in the database, such as exporter and importer names, type of packaging and insurance or freight costs, among others.

With this data source, we construct a monthly database for period 2002-2014 with the Brazilian dairy imports discriminated by country of origin (source). We focus on milk powder and cheese given their relevance in the total dairy trade, and the exporters are selected according to their market share.

Specifically, the model to be estimated includes the following dairy products: Whole Milk Powder (WMP), Skimmed Milk Powder (SMP) and (total) Cheese (Ch). We choose these products given that they represent approximately 80 percent of total value of Brazilian dairy imports. For WMP and SMP we take only Argentina ($a$) and Uruguay ($u$) as source countries because they capture on average more than 90 percent of milk powder imported by Brazil. For cheese, in addition to Argentina and Uruguay, we include the exporters grouped in 'others' ($o$), which are mostly European countries, although none individually maintains a particularly relevant and continuous participation in the time to be considered.

The response variables of the demand system are the budget-market shares, which are computed as the ratio between the expenditure in imports of a specific good from a certain source and the total expenditure in dairy imports (*i.e.* all imports included in the system); that is, $w_h^i = E_h^i / E$, where $E_h^i$ is the C.I.F. value of the imports of good $i$ from the country $h$ and $E = \sum_i \sum_h E_h^i$. Consequently, we have a complete demand system.

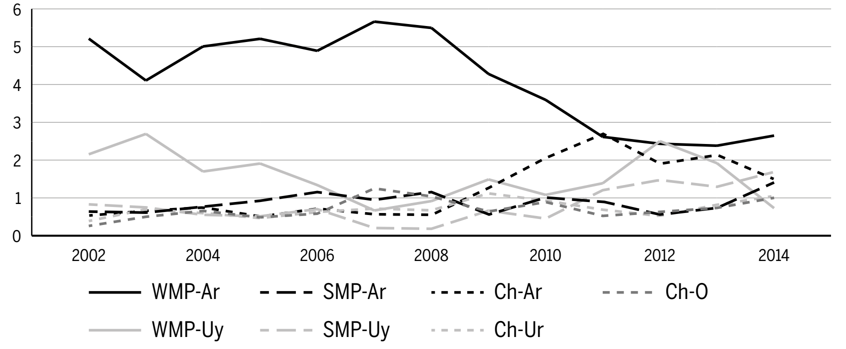
The demand prices of dairy imports are the C.I.F. prices, which are calculated as the ratio between the total C.I.F. value and the imported quantity (measured in net kilograms); that is, $p_h^i = E_h^i / q_h^i$.

Table 1 **Summary statistics of market shares and import prices for period 2002-2014**

| **Importer (*m*)** | **Product (*i*)** | **Exporter (*h*)** | $w_h^i$ | | $p_h^i$ | |
|---|---|---|---|---|---|---|
| | | | **Mean** | **SD** | **Mean** | **SD** |
| **Brazil** | WMP | Argentina | 0.412 | 0.163 | 3.074 | 1.083 |
| | WMP | Uruguay | 0.157 | 0.122 | 2.852 | 1.068 |
| | SMP | Argentina | 0.087 | 0.052 | 3.082 | 1.009 |
| | SMP | Uruguay | 0.08 | 0.082 | 2.99 | 1.15 |
| | Cheese | Argentina | 0.122 | 0.086 | 3.976 | 1.187 |
| | Cheese | Uruguay | 0.072 | 0.044 | 4.031 | 1.469 |
| | Cheese | Others | 0.071 | 0.043 | 6.423 | 2.327 |

*Source: Own calculations based on Custom Data (*Penta-*Transaction, 2014).*

Figure 1 **Evolution of yearly import market shares over the period 2002-2014**




*Source: Own calculations based on Custom Data (*Penta-*Transaction, 2014).*

In table 1 we present a summary of the statistics of the main variables involved in the system; that is, the shares and the import prices. From the average shares, we can observe the predominant positioning of Argentina in this market. In addition, in the Figure 1 we show the yearly evolution of these shares.

In this period the Argentine WMP represented on average more than 40 percent of the purchases have been decreasing, mainly since 2008. In contrast, Argentine cheese has a growing participation in the total Brazilian milk purchases. Although in the whole period they had an average share of 12.2 percent, since 2011 they exceeded 40 percent in some months. Uruguay's exports are concentrated in WMP. The gap with Argentina in SMP is considerably lower, with practically the same average shares. When comparing prices, we observe that the average prices of milk powder are higher in Argentina than in Uruguay. In contrast, a larger gap in cheese with the other exporting countries is found. This difference is mainly due to the composition of the types of cheeses traded.

## 3.2 **Econometric Model and Estimation**

Taking $= t = Jan - 2002,\ldots,Dec - 2014$ (T=156), the econometric or empirical version of the demand system (6) can be written as

$$w^{i}_{h,t} = \alpha^{i}_{h} + \sum_{j \in N} \sum_{k \in K_j} \gamma^{ij}_{hk} \ln p^{j}_{k,t} + \beta^{i}_{h} \ln\left(\frac{E_t}{\tilde{P}_t}\right) + \varepsilon^{i}_{h,t} \tag{18}$$

with $i \in N$, $h \in K_i$, where $\varepsilon^{i}_{h,t}$ is a random error term. For this particular case of the dairy import demand of Brazil, we have $N = \{WMP, SMP, Ch\}$, $K_{WMP} = \{a,u\}$, $K_{SMP} = \{a,u\}$ and $K_{Ch} = \{a,u,o\}$ using $a$, $u$ and $o$ to represent Argentina, Uruguay and the other exporting countries, respectively.

Given the potential interdependence of the system equations (18), it is generally assumed that there is a contemporary correlation between them (*i.e.* $Corr(\varepsilon^{i}_{h,,t}, \varepsilon^{j}_{k,t}) \neq 0$). For this reason, demand systems are usually estimated by *Zellner's* Seemingly Unrelated Equations (*SUR*) regressions via Maximum Likelihood (*ML*) or Feasible Generalized Least Squares (*FGLS*) using an estimator of the contemporary correlation matrix between the equations. This method is used in systems such as (18) and it is found in the early research of imports demand estimation (*e.g.* Seale Jr. *et al.*, 1992; Yang and Koo, 1994, Satyanarayana *et al.*, 1999), where generally yearly data with additional constraints (*e.g.* "block substitutability") are used to save degrees of freedom, even when they take a long period of time.

However, using time series can potentially create exogenous shocks that are related to preferences and prices that impact demand on a temporary or permanent basis (Holt and Goodwin, 2009). For the case of import demand, there are trade agreements, non-tariff barriers and other trade policies implemented by countries partners, and also changes in international food prices that are caused by shocks in world demand, that set a tendency or certain structural breaks that must be captured by the estimated demand. Additionally, when we work with quarterly or monthly data, as in the present case, it is likely that there is a seasonal component that causes displacements in the demand function. The usual way of introducing such variables or factors is through changes in the intercept (*e.g.* Seale Jr. *et al*, 2003, Feleke and Liu, 2005, Nahuelhual, 2005; Wan *et al*., 2010). In the case of introducing trend and seasonality variables, we will have a decomposition of the intercept in the form $\alpha^i_{h,t} = \alpha^i_{h,0} + \varsigma^i_h t_y + \sum_\tau \theta^i_{h,\tau} T_\tau$, where $T_\tau$ are dummy variables that reflect the seasonal *behavior* of the imports and $t_y$ is a yearly trend. Therefore, including trend and seasonality, (18) will be

$$w^i_{h,t} = \alpha^i_{h,0} + \sum_{j\in N}\sum_{k\in K_j} \gamma^{ij}_{hk} \ln p^j_{k,t} + \beta^i_h \ln\left(\frac{E_t}{\tilde{P}_t}\right) + \varsigma^i_h t_y + \sum_\tau \theta^i_{h,\tau} T_\tau + \varepsilon^i_{h,t} \qquad (19)$$

For this model, we have the following adding-up constraints:

$$\sum_i\sum_h \alpha^i_{h,0} = 1, \sum_i\sum_h \varsigma^i_h = 0, \sum_i\sum_h \theta^i_{h,\tau} = 0, \sum_i\sum_h \gamma^{ij}_{hk} = 0, \sum_i\sum_h \beta^i_h = 0,$$
$$\sum_i\sum_h \varsigma^i_h = 0, \sum_i\sum_h \theta^i_{h,\tau} = 0, \sum_i\sum_h \gamma^{ij}_{hk} = 0, \sum_i\sum_h \beta^i_h = 0$$

that are imposed to estimate (19) in addition to the homogeneity and symmetry, given by (9) and (10), respectively. For the estimation, one equation must be omitted (by singularity of error covariance matrix), while recovering the coefficients of the omitted equation from the adding-up, homogeneity and symmetry constraints.

For the construction of the seasonality variables, we analyze the mean of the budget shares $w^i_{h,t}$ for each month and computed along all years (2002-2014). We detect a seasonal behavior by trimester, mainly for WMP, revealing an opposite seasonal pattern between Argentina and Uruguay. Specifically, we define the following seasonality dummies variables: $T_1$ (taking as base) for the trimester corresponding to the months December-February, $T_2$ for March-May, $T_3$ para June-August and $T_4$ for September-November.

Although the problem of "zeros" in the working data set was avoided by choosing which exporting countries have a regular share in dairy imports, or by grouping exporters (*i.e.* the "Others") and dairy products (as cheeses), there are some months where the purchases to a certain origin are null. Although the proportion of zeros is negligible (in general not reaching 3 % of the data set), the computation of an implicit price (*CIF*) for these case results is impossible. Accordingly, the corresponding prices must be imputed to also consider the null imports as a possible corner solution from the utility maximization problem of the importer. In the literature it is very common to use linear regression to impute prices via the regression of the logarithm of price (to be imputed) on other variables (such as trend, other prices, quantities and the total outlay) that are complete in the sample, searching for a good fit for prediction purposes. Therefore, the logarithm of the import prices of good $i$ from country $h$ will be

$$\ln p^i_{h,t} = \begin{cases} \ln\left(E^i_{h,t} / q^i_{h,t}\right) & if \quad E^i_{h,t}, q^i_{h,t} \neq 0 \\ \mathbf{x}^T_t \hat{\mu}^i_h & otherwise \end{cases} \tag{20}$$

where $\mathbf{x}_t$ is the co-variables vector selected to predict the log-prices, and $\hat{\mu}^i_h$ is the parameter estimates ($OLS$) of the regression $E\left(\ln p^i_{h,t} \mid \mathbf{x}_t\right) = \mathbf{x}^T_t \mu^i_h$.

Given that $\tilde{P}_t$ in the Stone index includes their computation the market share $w^i_{h,t}$, which is the response variable in (18), there exists a potential endogeneity problem via simultaneity. Consequently, we need to adopt some correction of the Stone index to avoid this problem. Following Eales and Unnevehr (1988), we take the lagged shares in order to capture the changes in the market shares on the price index of imported dairy products. Therefore, the index price used in (17) will be

$$\ln \tilde{P}_t = \begin{cases} \sum_j \sum_k w^j_{k,1} \ln p^j_{k,1} & if \quad t = 1 \\ \sum_j \sum_k w^j_{k,t-1} \ln p^j_{k,t} & if \quad t > 1 \end{cases} \tag{21}$$

On the other hand, given that we will work with time series to estimate the demand system, it is convenient to model the serial correlation in the errors, additionally to contemporaneous correlation among equations. The existence of serial correlation in general is confirmed by the empirical investigations of demand that use time series (*e.g.* Washington and Kilmer, 2002; Seale Jr. *et al.*, 2003; Gohin and Féménia, 2009; and others). In particular, if $\varepsilon_t$ is a vector with $(n-1)$ errors terms of the system(19) of $n$ equations[4], then we can assume that this error term follows an autoregressive process of order 1; that is, $AR(1)$. So,

4 By singularity of the system, one equation must be omitted, so we have a dimension of $(n-1)$.

$$\varepsilon_t = \mathbf{R}\varepsilon_{t-1} + e_t \tag{22}$$

where $\mathbf{R}$ is an auto-correlation matrix of order $(n - 1) \times (n - 1)$ and $e_t$ is a vector of dimension $(n - 1) \times 1$ with $E(e_t) = \mathbf{0}$, $E(e_t e_t^T) = \Omega$ and $E(e_t e_t^T) = \mathbf{0}$ for all $t \neq s$. This enables both the contemporary correlation between equations (via $\Omega$) and the serial correlation for each equation (via $\mathbf{R}$) are considered. Then, following Holt (1998), the structure (22) implies a demand system of the form

$$\mathbf{w}_t = \psi\left(\mathbf{X}_t, \theta\right) + \mathbf{R}[\mathbf{w}_{t-1} - \psi\left(\mathbf{X}_{t-1}, \theta\right)] + e_t \tag{23}$$

where $\mathbf{w}_t$ is a vector with dimension $(n - 1)$, the elements are the shares $w_{h,t}^i$, $\mathbf{X}_t$ is the co-variables matrix (prices, expenditure, trend and dummies seasonal variables), $\theta$ is the parameter matrix of the system, that is $(\alpha_{h,0}^i, \gamma_{h,k}^{ij}, \beta_h^i, \theta_{h,\tau}^i, \varsigma_h^i)$ and $\psi$ (.) is the demand function corresponding to the *AIDS* system as in (19). From (23) we can observe that the parametrization of $\mathbf{R}$ must satisfy the adding-up restriction. There are several ways to parametrize $\mathbf{R}$ (*e.g.* Holt, 1998). In this paper we adopt a simple specification given by a single auto-correlation parameter following Berndt and Savin (1975). To estimate the system by adopting Berndt and Savin's (1975) method, we use the *R*-package 'erer' that was developed by Changyou Sun (2015).

Finally, it should be clarified that, in the context of demand analysis using time series, it is very common to use the dynamic specifications of demand systems. Some models introduce the dynamic by taking the involved variables in differences (*e.g.* Seale Jr. *et al.*, 2003, Muhammad *et al.* 2004; 2007; Harri *et al.* 2010), or using as predictor the sum of the lagged response variables as in Feleke and Liu (2005). A vast empirical literature (*e.g.* Ben Kaabia and Gil, 2007; Nzaku and Huston, 2009, Wan *et al.*, 2010; Nzaku *et al.*, 2012, among others) found unit roots in the series that confirm non-stationarity and cointegration of the data, which they used to justify the use of an error correction model specification for the demand system. However, as Holt and Goodwin (2010) argue, this dynamic approach has an important limitation because the response variable (*i.e.* the budget/market share) is by definition bounded in [0,1], which is inconsistent with unit root behavior and, therefore, with the use of this variable in first-differences. Even for the empirical case discussed here, we corroborate through the application of unit root test (specifically, Augmented

Dickey-Fuller and Phillips-Perron test) that the shares are stationary. In turn, prices can be stationary by imposing homogeneity via their expression in terms of relative prices (Gil *et al*., 2004). Therefore, in the present paper we decide to work with static systems, thereby limiting the results to obtaining long-term elasticities.

## 4 **Results**

In table 2 we present the estimated coefficients of the Brazilian demand for dairy imports differentiated by countries of origin (sources). We exclude the equation of cheese imported from the others (*i.e.* $w_o^{Ch}$) by the singularity of the system, but it can be recovered from homogeneity, adding-up, and symmetry constraints.

First, seasonal variables are shown to be statistically significant only for the case of WMP from Argentina, indicating a smaller share in the Brazilian market in all of the trimesters compared to the months of December to February ($T_1$). This reduction in the Argentinian share seems to be compensated by a higher Uruguayan share, where the seasonality is significant only in $T_2$, but with coefficients contrary to those of Argentina in all trimesters. For the other dairy products (SMP and cheese), there is no relevant seasonal behavior in market shares.

The trend variable shows that the share of WMP in the total expenditure on dairy imports has been decreasing for both sources, in favor of cheese mainly of Argentinian origin. These trends respond more to an increase in imported quantities of cheese than a reduction in WMP quantities. Up to 2014, increased Brazilian purchasing power and changes in diets incorporating more added value products like cheese could help to explain its higher demand (Freitas, 2012). For the SMP, although slight, the share in total expenditure on dairy imports has also been increasing.

The coefficients $\beta_h^i$ for milk powder from all countries are statistically significant at the 1 percent level, indicating that changes in total real expenditure have a significant impact on the budget shares of milk powder, being negative on WMP but positive on SMP. That is, an increase (decrease) in Brazilian total real expenditure in dairy imports is associated with a fall (increase) in the WMP's share and an increase (fall) in the SMP's share, with a higher effect for imports that come from Uruguay. Conversely, for Argentine cheese, the changes in total expenditure do not show a significant effect on its budget share, but for Uruguayan cheese there is a significant negative association at 5 percent level. For the price coefficients, we observe that the prices of the cheese are significant, particularly in the Argentine and Uruguayan cheese share equations.

As can be expected of this type of microeconometric model, we observe an acceptable goodness of fit in terms of the R-square, mainly for WMP and Argentine cheese. From the diagnostic tests, the Breush-Godfrey test (test B-G) shows that the results of the correction for serial correlation with a structure AR(1) are effective, not rejecting the null hypothesis in five of the six equations[5]. From the Breusch-Pagan test, at the 5 percent level, we find that in four of the six equations the hypothesis of homoscedastic errors cannot be rejected at 1 percent, in five of the six demand equations the homoscedasticity assumption is not rejected. The Ramsey's Regression Specification Error Test (RESET) shows that at a 1 percent level, in five of the six equations there is insufficient statistical evidence that there is a functional misspecification. Therefore, with the exception of the share equation for Argentine cheese, the estimated AIDS of Brazilian dairy imports shows a good statistical performance.

We estimate the demand system imposing the constraints (8)-(10) for its consistency with economic theory. Although this theoretical requirement must be maintained in coherence with the microeconomic behavior, testing these constraints empirically is an usual practice and is also recommended (Latinen, 1978; Meisner, 1979). Basically it is done to check if the observed and adjusted data are consistent with the behavior predicted by the consumer theory in an import demand framework. Table 3 presents the likelihood ratio tests for these theoretical restrictions. In addition to the homogeneity and symmetry constraints, we test the serial correlation modeled by an AR(1). The likelihood ratio test, rejects the non-existence of first-order serial correlation, which is consistent with the Breusch-Godfrey test. In addition, it can be observed that no economic restriction is rejected. Therefore, we can conclude that in this case the theoretical constraints of the demand system are compatible with the empirical results of the model.

5 When the system is estimated via SUR regression without an AR(1) correction, the results of B-G test do not reject the serial correlation in all equations.

Table 2 **Demand System Coefficients of Brazilian Dairy Imports**

| Variables | $w_a^{WMP}$ | $w_u^{WMP}$ | $w_a^{SMP}$ | $w_u^{SMP}$ | $w_a^{Ch}$ | $w_u^{Ch}$ |
|---|---|---|---|---|---|---|
| $T_2$ | −0.087*** (0.029) | 0.052** (0.026) | −0.004 (0.012) | 0.024 (0.019) | −0.013 (0.014) | −0.002 (0.011) |
| $T_3$ | −0.087*** (0.030) | 0.019 (0.027) | 0.0001 (0.008) | 0.015 (0.019) | 0.008 (0.014) | 0.015 (0.010) |
| $T_4$ | −0.051* (0.028) | 0.003 (0.022) | 0.002 (0.017) | 0.008 (0.018) | 0.009 (0.014) | 0.022** (0.009) |
| *trend* | −0.022*** (0.004) | −0.010*** (0.003) | 0.004** (0.002) | 0.003 (0.002) | 0.014*** (0.002) | 0.004*** (0.001) |
| $\ln(E/P)$ | −0.060*** (0.019) | 0.100*** (0.017) | −0.026*** (0.009) | 0.040*** (0.013) | 0.008 (0.009) | −0.015** (0.006) |
| $\ln p_a^{WMP}$ | −0.084 (0.172) | −0.135 (0.087) | −0.016 (0.07) | 0.119 (0.092) | −0.002 (0.067) | 0.086** (0.041) |
| $\ln p_u^{WMP}$ | −0.135 (0.087) | 0.101 (0.08) | 0.045 (0.039) | −0.005 (0.051) | 0.052 (0.036) | 0.014 (0.024) |
| $\ln p_a^{SMP}$ | −0.016 (0.07) | 0.045 (0.039) | 0.018 (0.058) | −0.042 (0.048) | 0.025 (0.030) | −0.041* (0.022) |
| $\ln p_u^{SMP}$ | 0.119 (0.092) | −0.005 (0.051) | −0.042 (0.048) | −0.001 (0.083) | −0.056 (0.038) | −0.047* (0.028) |
| $\ln p_a^{Ch}$ | −0.002 (0.067) | 0.052 (0.036) | 0.025 (0.030) | −0.056 (0.038) | −0.069* (0.035) | 0.058*** (0.019) |
| $\ln p_u^{Ch}$ | 0.086** (0.041) | 0.014 (0.024) | −0.041* (0.022) | −0.047* (0.028) | 0.058*** (0.019) | −0.075*** (0.019) |
| $\ln p_o^{Ch}$ | 0.034 (0.034) | −0.071*** (0.021) | 0.012 (0.019) | 0.033 (0.023) | −0.008 (0.016) | 0.005 (0.012) |
| $\alpha_j$ | 1.503*** (0.281) | −1.290*** (0.252) | 0.459*** (0.128) | −0.561*** (0.186) | −0.093 (0.134) | 0.268*** (0.092) |
| $R^2$ | 0.397 | 0.254 | 0.1 | 0.174 | 0.452 | 0.214 |
| B-G Test † | 0.21 [0.65] | 0.018 [0.89] | 0.07 [0.79] | 0.472 [0.49] | 17.531 [0.00] | 0.553 [0.46] |
| B-G Test ‡ | 31.761 [0.00] | 16.11 [0.19] | 10.275 [0.59] | 15.719 [0.2] | 23.796 [0.02] | 12.382 [0.42] |
| RESET ⁑ | 3.778 [0.03] | 0.272 [0.76] | 2.416 [0.09] | 0.785 [0.46] | 10.432 [0.00] | 2.058 [0.13] |

*Sou Note: Standard errors in parentheses. ***p <0.01%; **p <0.05%; *p <0.1%. P-values in brackets.*
*† H0: no serial correlation. ‡ H0: no heteroscedasticity. ⁑ H0: no functional misspecification error.*

Table 3 **LR-Tests: Autocorrelation, Homogeneity and Symmetry**

| | Log-Likelihood ($\ell$) | $-2[\ell(\theta^{R})-\ell(\theta^{NoR})]$ | df |
|---|---|---|---|
| **No –AR(1)** | 1499.1 | | |
| **AR(1)** | 1529.9 | 61.6 | (1) |
| **Homogeneity** | 1525.7 | 8.4158 | (6) |
| **Symmetry** | 1520.5 | 18.692 | (15) |
| **Homogeneity & Symmetry** | 1518.9 | 21.941 | (21) |

In table 4 we present the expenditure and own-price elasticities, as computed from (12), (16) and (17), evaluated at the average shares, with the standard errors obtained via delta method.

We found that the Brazilian imports demand for Uruguayan WMP and SMP and for Argentine cheese are elastic with respect to the total expenditure in dairy imports, with values equal to 1.6, 1.51 and 1.18, respectively. In contrast, the expenditure elasticities are less than the unit for Argentine WMP and SMP (0.8 and 0.76, respectively), and also for cheese from Uruguay and other sources (0.847 and 0.387, respectively). This implies that if Brazilian dairy imports expenditure increases by 1 percent, then the quantity of Uruguayan milk powder and Argentine cheese will increases by more than 1 percent, while the imported quantities of Argentine milk powder and cheese from Argentina's competing countries would increase by a smaller percentage. In terms of the (micro) economic categorization of goods, these results would reveal that, although for Brazil all dairy imports are “normal goods”, Argentine cheese and Uruguayan milk powder would constitute “luxury goods”; once we assume that total dairy expenditure constitutes a *proxy* for income the others are “necessities”. When we performed a Wald z-test to contrast equality of elasticities, we found that there is no significant difference between WMP and SMP from the same exporter. Nevertheless, there exist statistical difference in the expenditure elasticities between the dairy imports from Argentina and Uruguay. More precisely, the import expenditures on the Uruguayan milk powder are statistically more expenditure-elastic than that of Argentine (with the statistic values $z = 7.87$ for WMP and $z = 4.497$ for SMP). The import quota for powdered milk from Argentina imposed since 2009, could help to explain this significant difference between expenditure elasticities: In this way, a higher import expenditure on powdered milk would necessarily mean a more than proportional increase in those from its competitor Uruguay.

All own-price elasticities are negative and significant, except for the WMP and SMP from Uruguay which are not statistically significant. In all cases, uncompensated (Marshallian) price elasticities are greater in absolute value than the compensated (Hicksian) price elasticities. This means that the "income effects" are positive or, more precisely, the effects of changes in the price on the total expenditure on imported dairy products. Specifically, the gap between compensated and non-compensated elasticities are higher for those goods with a larger share of total expenditure (such as Argentine WMP and cheese) because the effect of changes in the prices of such products has a greater weight on actual expenditure. In particular, the Argentina WMP is elastic when the uncompensated elasticity is considered while it is inelastic according to the compensated elasticity. More precisely, a percentage reduction in the price of the Argentine WMP increases its quantity demand in a greater proportion if the income effect is considered. However, if actual expenditure on milk imports remains constant, then a 10 percent increase in the price of the Argentine WMP would increase its quantity demand by 8 percent. Nevertheless, if we take into account the variability of both elasticities (compensated and not compensated) for this product (WMP), the difference is not statistically significant (the z-test gives a statistic equal to 0.569). In fact, both are not statistically different from one, concluding that on average their demand elasticities could be considered unitary. In contrast, for Argentine SMP and cheese, even if the real expenditure remains constant, the demands are elastic with respect to their own price. But as in the previous case, the null hypothesis of unit elasticities cannot be rejected. The greatest elasticity (in absolute value) is revealed for Uruguayan cheese, showing that an increase (decrease) of 10 percent in its price, decreases (increases) their quantity demand by more than 20 percent. In turn, for cheese from other origins, the elasticity is practically unitary.

These results of the own elasticities, would show in general, that the Brazilian import demand is very sensitive to changes in prices. These results are consistent with the price-elasticities estimated by Santos and Barros (2006) who used an empirical import equation with times series estimation for the period 1991-2003. For our estimation, a possible reason for this import demand characteristic is the growing domestic production revealed in the period 2002-2014. It could have helped to substitute some of imported dairy products (Giovinazzo Spers *et al.*, 2013; Videla *et al.*, 2017), raising the demand elasticity for imported dairy products respect to their own prices. In terms of some policies effects, any commercial policy or changes in transport costs that translate into price increases imply a reduction in the demand for each imported dairy product in equal or greater proportion to such price changes. The exception of this is the WMP from Uruguay.

Table 4 **Expenditure ($\eta^{E}_{i,h}$) and Own-Price Elasticities ($\eta^{ii}_{hhU}$ & $\eta^{ii}_{hhC}$)**

| Product | Exporter | Expenditure Elasticity | Own-Price Elasticity | |
|---|---|---|---|---|
| | | | Marshallian | Hicksian |
| WMP | Argentina | 0.800***<br>(0.041) | −1.131**<br>(0.407) | −0.802*<br>(0.408) |
| WMP | Uruguay | 1.600***<br>(0.093) | 0.074<br>(0.481) | 0.325<br>(0.479) |
| SMP | Argentina | 0.761***<br>(0.090) | −1.580**<br>(0.676) | −1.514**<br>(0.674) |
| SMP | Uruguay | 1.517***<br>(0.142) | −1.321<br>(1.007) | −1.199<br>(1.011) |
| Cheese | Argentina | 1.182***<br>(0.070) | −1.434***<br>(0.274) | −1.290***<br>(0.275) |
| Cheese | Uruguay | 0.847***<br>(0.078) | −2.203***<br>(0.238) | −2.143***<br>(0.237) |
| Cheese | Others | 0.387***<br>(0.063) | −0.988***<br>(0.185) | −0.961***<br>(0.185) |

*Note: S.E. in parentheses. ***p <0.01%; **p <0.05%; *p <0.1%.*

The difference between the Argentine and Uruguayan WMP price elasticities is remarkable, being almost zero for the latter case. This could be explained by the weight of each exporter in the Brazilian market. However, for the selected period, and particularly since 2008 (coinciding with a period of strong government intervention on dairy Argentine exports via trade taxes and restrictions and the import quota imposed by Brazil), there is at the same time an increasing trend in the quantities imported from Uruguay and a decreasing trend in the quantities imported from Argentina, closing the gap between the market shares of both countries for this dairy product. While the Uruguayan WMP managed to maintain its share at around 15 percent, Argentine WMP strongly decreased its share, going from an average of 50 percent before 2009 to 27.3 percent since 2010. On the one hand, the greater elasticity of Argentina with respect to Uruguay would show the loss of competitiveness of the Argentine WMP in the Brazilian market. Furthermore, the lower price elasticity in Argentine cheese compared to Uruguayan cheese would reveal the greater competitive potential of the Argentine cheese.

Table 5 **Cross-Price Elasticities of Brazilian Dairy Import Market**

| Quantity | Price | | | | | | |
|---|---|---|---|---|---|---|---|
| | WMP-Ar | WMP-Uy | SMP-Ar | SMP-Uy | Ch-Ar | Ch-Uy | Ch-O |
| | **Marshallian** | | | | | | |
| WMP-Ar | – | −0.494** | 0.077 | 0.370* | −0.059 | 0.273*** | 0.164** |
| | – | (0.202) | (0.175) | (0.218) | (0.135) | (0.094) | (0.082) |
| WMP-Uy | −1.628*** | – | 0.112 | −0.17 | 0.513** | 0.03 | −0.530*** |
| | (0.528) | – | (0.238) | (0.323) | (0.220) | (0.152) | (0.119) |
| SMP-Ar | 0.378 | 0.331 | – | −0.015 | 0.411 | −0.463* | 0.178 |
| | (0.818) | (0.427) | – | (0.556) | (0.327) | (0.243) | (0.226) |
| SMP-Uy | 1.602 | −0.319 | −0.082 | – | −0.985** | −0.676** | 0.265 |
| | (1.128) | (0.638) | (0.603) | – | (0.450) | (0.329) | (0.292) |
| Ch-Ar | −0.356 | 0.726** | 0.259 | −0.623** | – | 0.466*** | −0.220* |
| | (0.458) | (0.285) | (0.237) | (0.298) | – | (0.146) | (0.127) |
| Ch-Uy | 1.553*** | 0.183 | −0.573* | −0.704* | 0.834*** | – | 0.065 |
| | (0.54) | (0.33) | (0.296) | (0.365) | (0.245) | – | (0.151) |
| Ch-O | 1.124** | −0.983*** | 0.252 | 0.391 | −0.281 | 0.098 | – |
| | (0.471) | (0.276) | (0.276) | (0.329) | (0.216) | (0.150) | – |
| | **Hicksian** | | | | | | |
| WMP-Ar | – | −0.369* | 0.147 | 0.434* | 0.039 | 0.330*** | 0.221** |
| | – | (0.202) | (0.174) | (0.219) | (0.136) | (0.094) | (0.081) |
| WMP-Uy | −0.969* | – | 0.251 | −0.042 | 0.708*** | 0.144 | −0.417*** |
| | (0.530) | – | (0.237) | (0.326) | (0.221) | (0.150) | (0.124) |
| SMP-Ar | 0.691 | 0.45 | – | 0.046 | 0.504 | −0.409* | 0.231 |
| | (0.820) | (0.424) | – | (0.554) | (0.329) | (0.242) | (0.223) |
| SMP-Uy | 2.227* | −0.082 | 0.05 | – | −0.801* | −0.567* | 0.372 |
| | (1.122) | (0.636) | (0.602) | – | (0.453) | (0.327) | (0.289) |
| Ch-Ar | 0.131 | 0.911*** | 0.362 | −0.528* | – | 0.551*** | −0.136 |
| | (0.456) | (0.284) | (0.236) | (0.299) | – | (0.145) | (0.125) |
| Ch-Uy | 1.901*** | 0.316 | −0.499* | −0.636* | 0.937*** | – | 0.125 |
| | (0.541) | (0.329) | (0.295) | (0.367) | (0.246) | – | (0.15) |
| Ch-O | 1.284** | −0.922*** | 0.286 | 0.422 | −0.234 | 0.126 | – |
| | (0.471) | (0.275) | (0.276) | (0.328) | (0.216) | (0.151) | – |

*Note: S.E. in parentheses. ***p <0.01%; **p <0.05%; *p <0.1%.*

Table 5 shows the cross-price elasticities derived from demand system. Although cross-elasticized Marshallian elasticities are also presented for exposition purposes, to analyze the degree of substitution or competition among the different imported dairy products, the Hicksian elasticities are more informative because they represent pure substitution effects (*i.e.* net of the income effect).

In general terms, the cross-price elasticities show the competitive relationship among different products and exporters. If they are positive, then the goods are substitutes, being complementary otherwise. In the import demand studies, it is expected that the same product but with different origins are substitute but this does not necessarily occur in the empirical findings. In fact, parallel movements in the exchange rates between two exporters, as well as the restrictions imposed on the demand systems (homogeneity and symmetry), may contribute to the apparent relations of complementarity (Yang and Koo, 1994: 406). For example, in Table 5 we observe negative cross-price elasticities between the Argentina and Uruguayan WMP. This means that an increase in the price of the Argentine (Uruguayan) WMP reduces the quantity demand of Uruguayan (Argentine) WMP.

This can be explained because the WMP is a commodity with a low degree of differentiation, and Argentina and Uruguay are neighbouring countries (and, therefore, with similar transport costs). Therefore, it is very probable that price movements will be parallel, which would be captured through a negative cross-price elasticity. In other cases, such as between Uruguayan (Argentine) SMP and Argentine (Uruguayan) cheeses, the negative sign is more difficult to explain and interpret. Other complementarities, such as different goods from the same country of origin (such as Uruguayan SMP and cheese), have more economic sense through the transport economies and international negotiations that an importing country has when buying different products of a certain exporter.

Focusing on the degree of substitution between the imported products, we can observe that between the Argentine WMP and the Uruguayan SMP there is a significant cross-effect. More precisely, the effect of the Argentine WMP prices on the demand for Uruguayan SMP (2.227) is greater than the effect of changes in the prices of the Uruguayan SMP on the quantity demand for Argentine WMP (0.434). Specifically, a decrease in the price of the Argentine WMP would lead to a reduction in the demand for Uruguayan SMP in a proportion that duplicates the price change (*i.e.* $\eta_{a,u}^{WMP,SMP} = 2.227$). This cross-effect is greater than the own-effect, which could be explained by the share of the Argentina WMP in the total dairy imported by Brazil. There is also a significant and high substitutability between Argentine WMP and cheese from competing countries (*i.e.* Uruguay and Others). Meanwhile, the cross-price elasticity is high and statistically significant between Argentine cheese and the Uruguayan WMP ($\eta_{a,u}^{Ch,WMP} = 0.911$ and $\eta_{u,a}^{WMP,Ch} = 0.708$).

The cross-price elasticities show an important level of competition between Argentine and Uruguayan cheese, revealing an asymmetric substitution response. An increase of 10 percent in the average price of Argentine cheese would increase the quantity demand of Uruguayan cheese by almost 10 percent, while if Uruguayan cheese is increased by 10 percent, then the substitution for Argentine cheese would be proportionally lower (specifically, 5.51 percent).

## 5 **Conclusions**

Based on the AIDS approach, the present paper estimates Brazilian import demand for dairy products. For the period that we studied, the results show that Argentina and Uruguay maintain a majority share and Brazil constitutes one of the most important destinations for their dairy exports. That is not surprising given their neighbourhood and Mercosul preferential treatment agreement.

In terms of the sensitivity of Brazilian demand to changes in prices, Argentine cheese seems to have a greater competitive potential than WMP, despite the high participation of the latter in the market. However, the price substitutability facing Argentina is considerable. The opposite occurs for Uruguayan WMP, which has gained market share during the last few years. In this way, the knowledge of the Brazilian demand response also helps to understand the competitiveness of the dairy sectors of these Mercosul countries.

Based on the values of the estimated expenditure elasticities, we find that for the Brazilian market, a higher expenditure on dairy imports is distributed in a way that increases in a greater proportion the quantity demand of Uruguayan milk powder and Argentine cheese. Meanwhile, Argentine milk powder and cheese from its competitors increases in a smaller proportion.

As the issue of Brazilian dairy import demand elasticities with emphasis in the Mercosul context has been underexplored, this paper intends to contribute to fill the gap. However, a following up with a more disaggregated dairy model, more detailed export countries specifications or/and the inclusion of domestic production in the demand system jointly with imported products would be helpful to reinforce the findings and achieve a better understanding of the results.

We expect the above information could be used by sector participants and producers organizations in their production and commercial strategies. As dairy trade has been subject to controversies in the past, it could also help government decision makers to delineate trade measures and make projections for their potential implementation. Not less important, in a context in which free trade agreements are being negotiated (such as with the European Union), knowledge of the elasticities may provide additional elements to evaluate the behavior of flow trades with other countries, using the Mercosur experience.

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