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Determinants of Brazilian agribusiness exports to China

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Abstract

Brazil is a major global agribusiness trader, a leading producer and exporter of grain, meat and sugar while China is notable as a crucial trading partner for Brazilian agribusiness. With these aspects in mind, this article sets out to analyze the trade flow determinants of agribusiness from Brazilian states to China. To do so, foreign income and real exchange rate variables were constructed, using state-specific trade weights. The estimates were calculated using static and dynamic panel data. The results suggest that exports from Brazilian states to China are elastic in terms of foreign income and inelastic in terms of the variable production level of the states. Price (exchange and international price) did not rank as a determinant.

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1. Introduction

The solid economic performance of Brazilian agribusiness is widely recognized. According to the Center for Advanced Studies in Applied Economics (CEPEA, 2014), the agribusiness Gross Domestic Product (GDP) in 2013 yielded R$1,099.4 million at 2013 prices. This amount accounted for more than 20% of total national GDP.

In terms of agribusiness exports, the balance of trade is very positive. In 2013 total exports reached a record US$101.5 billion, an increase of 4% over the previous year, which had also been a record (CEPEA, 2014). This expansion was due to volume, which reached a new historical high, with an increase of 14.2%, as average export prices in dollars fell 7.5% over the period. Considering that the 2013 balance of trade showed a small surplus of $2.5 billion, agribusiness contributed to easing the trade balance of other sectors of the economy, as it generated a trade surplus of nearly US$83 billion.

The segment which registered highest exports in 2013 was the soy complex, yielding a total of approximately US$30.1 billion. This was followed by the meat complex with US$16.8 billion and in third place, the sucro-alcohol complex, with US$13.7 billion (MAPA, 2014)[1].

On the question of the destination for Latin American and Brazilian exports, China ranks as a very strong partner (see Jenkins *et al.*,2008; Feistel and Hidalgo, 2012; Velloria, 2012). In 2013, Brazilian agribusiness exports to China amounted more than US$45.7 billion (MAPA, 2014). From 1997 to 2013 the annual growth rate of exports to China was a very significant 29.94%. Total exports in 2013 hit US$66 billion which included exports to the European Union (EU) (28 countries), United States, Japan and the Middle East. In the analyzed period the China's economy grown up around 4.5 times, and also China's importation from the world increased significantly, being almost 350% in the period. In this scenario, the China´s economy became an important global player and the Brazilian exportation to China has grew in the same rhythm.

This article contributes to identify the determinants of agribusiness exports from Brazilian states to China, their main partner. More specifically, it set out to present an economic model which would support an econometric panel data model specification capable of identifying and quantifying the variables which affect agribusiness exports. This research breaks new ground in that it emphasizes the role of state-specific, trade-weighted foreign income and real exchange rates as determinants of the bilateral trade flows of Brazil-China agribusiness, this strategy is motivated by paper of Cronovich and Gazel (1998). For estimates and tests, the dynamic panel data procedure is used. It allows for the control of potential endogeneity and reverse causality. Thus, it is expected that by presenting new results this research will contribute to discussion on the topic and assist in drawing up policies for Brazil, which export a large amount of soybeans and soybean oil to China.

After this introduction, the article goes on to present a brief overview of Brazilian agribusiness exports. The third section deals with related empirical literature while the following presents data and empirical strategy. The penultimate section treats of estimates and analysis of the results and the study ends with concluding remarks.

2. Brazilian agribusiness exports to China

The relevance of agribusiness for the Brazilian GDP cannot be denied. In 2012, in current values of that year, it reached R$1,099.4 billion. Agricultural agribusiness was responsible for R$766.8 billion of this and the remainder, R$332.6 billion, came from livestock.

[1] Foreign Trade Statistics of Brazilian Agribusiness – AGROSTAT from Ministry Agriculture.

About exports, it accounted for approximately R$252 billion. This means it is a key and vibrant sector of the Brazilian economy (CEPEA, 2014).

In addition to the evolution of growth rate, it is also important to understand the relative participation of the states in the volume exported. During the period under consideration, China is undoubtedly the main market[2], consuming approximately one third of Brazilian agribusiness exports.

Figure 1 shows the development of agribusiness exports to the Chinese market in the period under study (2003-2011)[3]. In 2003, around 9.83% of Brazilian agribusiness exports went to China, while in 2011 this market share jumped to approximately 26.11%, a significant increase of 16.28 percentage points in the period. In monetary terms (US$) Brazil exported 2.94 billion to China in 2003, and this figure jumped to 16.51 billion in 2003/2011; therefore, agribusiness exports to China increased more than five times while total exports increased by around 2.26 times in the same period.

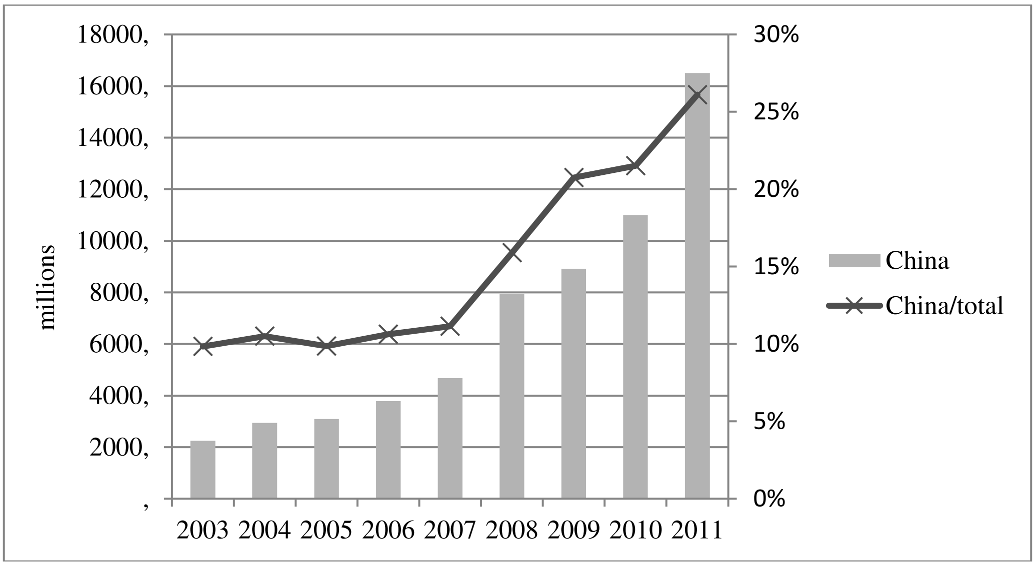

Figure 1 – Evolution of Brazilian agribusiness exports (US$) to the Chinese market: 2003-2011.
Source: Drawn up by the authors using MAPA-AGROSTAT database (2014).

Figure 1 shows the overall evolution of Brazilian agribusiness exports, however, the distribution of exports between the states is very diverse. In some states there was a significant variation in exports while in others this variation was below average, or even negative.

Figure 2 shows the distribution of variation in agribusiness exports by Brazilian states[4] between 2003 and 2011. It is noteworthy that the states which export more agribusiness products (São Paulo, Paraná, Rio Grande do Sul and Mato Grosso) were not those with greater variation. This is due to the fact that because the level of exports is very high any variation will apply on a large comparison base (see Table 3, Annex A). The greatest variation was recorded in the Federal District which exports relatively little. This was followed by Tocantins state which is not a traditional trader either. In 2011 the state of São Paulo yielded 19.7% of the exports, followed by Paraná with 13.6%, and Rio Grande do Sul with 12.5%. The data show that there is a certain concentration on the part of exporting states.

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[2] Others important markets see Almeida et al. (2012).
[3] The reporting period ends in 2011 due to availability of part of the survey data (see Section 4.1).
[4] Four states were excluded from the sample because of zero value for exports in the period under study (see Section 4.1).

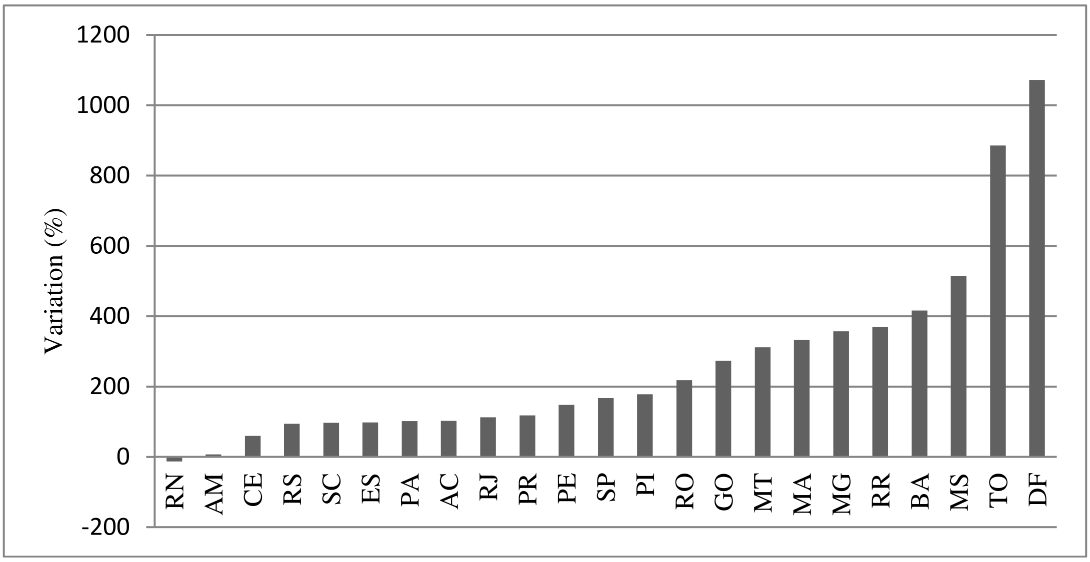

Figure 2 – Variation in exports between 2003 and 2011 in the brazilian States.
Source: Drawn up by the authors using MAPA-AGROSTAT database (2014).

On the agenda of agribusiness exports in 2013, the soy complex stands out, reaching nearly US$31 billion, followed by meat and sucro-alcohol, US$16.8 and 13.7 billion, respectively (MAPA, 2014).

3. Related empirical literature

With the considerable growth in trade between Brazil and China, the debate on the determinants of trade has intensified. In relation to agribusiness, there are several studies on the theme; however, estimates are based on the aggregate time-series procedure (see Mortatti *et al*, 2011). Thus, new research which takes the heterogeneity of the Brazilian states into consideration in the estimates could lead to more reliable results, as certain states export a considerable amount while others export much less.

Ferreira *et al*. (2006) analyzed the impact of the 1999 exchange rate variation on the trade balance of Brazilian agribusiness. They posit that the system of exchange bands exerted a negative influence on agribusiness exports in the first five years of the Real Plan. From 1999 onwards, results indicated that the floating exchange rates positively affected exports. Neves and Lélis (2007) estimated elasticities of the aggregated exports of Brazilian states and the results confirmed that foreign exchange and income had an inelastic effect on exports for the 1992-2004 period under analysis.

Barros and Silva (2008) set out to analyze the contribution of Brazilian agribusiness to the trade balance for the 1989-2005 period, using a new classification for agribusiness. They developed specific models for exports and imports and observed that a 1% increase in attractiveness - given by the product of exchange and foreign prices – immediately boosts exports of non-processed agricultural products by 1.71% and stabilizes at 2% after a few quarters. Another significant result was that a 1% increase in GDP has the impact of cutting exports of agricultural products by 1.7%.

Stocly *et al*. (2011) analyzed the determinants of exports and imports of Brazilian agribusiness from 1995 to 2009. They concluded that exports grew at an average rate of 4.63% per year, while imports decreased at an average annual rate of 4.46%. With regard to the determinants of exports, for them the real effective exchange rate was not the variable responsible. As for imports, this variable affected them negatively. Devadoss et al. (2014) present the negative impact that Yuan exchange rate devaluation would have on United States exports.

Using the time series procedure, Mortatti *et al*. (2011) identified the Brazil-China trade determinants and the results indicated that, in the case of commodities, exports proved elastic in relation to variations in income and inelastic in relation to exchange rates and prices. On analyzing the trade flow between Brazil and China and focusing on an analysis of comparative advantages, Feistel and Hidalgo (2012) found that trade between these two countries could be compatible with the principles of comparative advantage, assuming that Brazil is relatively abundant in natural resources and relatively scarce in capital.

4. Data and empirical strategy
4.1. Concept and description of variables

The data used in this study refer to twenty-two Brazilian states[5] and the Federal District from 2003 to 2011. This research breaks new ground by emphasizing the role of state-specific, trade-weighted real exchange rates and foreign income as determinants of the bilateral trade flow for Brazil-China agribusiness over the 2003 to 2011 period[6]. The construction of the state-specific trade weight is represented by the trade share of each state i, for a given country j, in this case China. The construction of the other variables follows the proposal of Cronovich and Gazel (1998).

Thus, the share of state i's exports to country j for year t can be presented as:

$$Weight(w)_{i,j,t} = X_{i,j,t} \Big/ \sum_{k} X_{i,j,t} \tag{1}$$

in which $X_{i,j,t}$ are state i exports to country j (China) in year t. Data on the agribusiness exports of each state were obtained through the AGROSTAT system[7] of the Brazilian Ministry of Agriculture (MAPA).

To construct the trade-weighted GDP (Y) of each state's trade partners the GDP (PPP, constant 2005 US$) of the World Bank is used. Considering $Y^*_{j,t}$ the real GDP of country j for year t, then the trade-weighted foreign GDP* for state i in year t is represented by:

$$wy^*_{j,t} = w_{i,j,t} Y^*_{j,t} \tag{2}$$

State i's trade-weighted real exchange rate was obtained from data on the nominal exchange rate denoted by $E_{R\$/\$,t}$ and the consumer price index (CPI) of each country - P_j for the foreign price index and P_i for the Brazilian. Thus, the trade-weighted real exchange rate between the state i and the foreign country j (China) in each period (t) is represented by the following expression:

[5] Alagoas (AL), Amapá (AP), Paraíba (PB) and Sergipe (SE) states were not considered in the analysis as their export flow to China was equal to zero for most of the period under study.
[6] This period began in 2003 as it was then that China began to import agricultural commodities from all the States considered here. It ends in 2011 as all the data (Chinese GDP and world agricultural commodities prices) were not available after this date.
[7] Available at: http://sistemasweb.agricultura.gov.br/pages/AGROSTAT.html.

$$w\varepsilon_{i,j,t} = w_{i,j,t}\varepsilon_{i,j,t} \qquad (3)$$

in which, $\varepsilon_{R\$,j,t} = E_{R\$/j,t} P_{j,t} / P_{i,t}$ is the real exchange rate between Brazil and the country j (China). Data on the nominal exchange rate and foreign price index were obtained from the International Monetary Fund (IMF); data on the Brazilian price index and GDP of the respective states were obtained from the Institute for Applied Economic Research (IPEA). In addition, the commodity price index (CP) was obtained from the World Bank[8].

Descriptions of the variables used are:

i. Exports, *x*: refers to agribusiness exports of each state obtained using the MAPA (by AGROSTAT system);

ii. Foreign GDP, *wy**: is China's GDP (purchasing power parity-PPP, constant 2005 US$) obtained from the World Bank;

iii. Real exchange rate, *wε*: the real exchange rate was calculated from the nominal Brazil-China cross exchange rates. The Brazilian nominal exchange rate (R$/US$) and the Chinese nominal exchange rate (yuan/US$) and price index (CPI) were obtained from the International Monetary Fund. The Brazilian consumer price index (CPI) was obtained from IPEA;

iv. GDP, *y*: the GDP of each state (in R$ for 2000) was obtained from IPEA;

v. Commodity prices, *cp*: the international price index of agricultural commodities provided by the World Bank Global Economic Monitor Commodities;

vi. Weight *w*: weight variable in each state's share of agribusiness exports to China in terms of total state agribusiness exports, according to equation (1).

Brazil's continental dimensions are reflected in the heterogeneity of the data of Brazilian states' exports and in such a context the panel data procedure to be presented in the next section is considered the most appropriate.

4.2. The panel data

To achieve the empirical objectives of the study a panel data will be set up and both the static econometric procedure and that of dynamic equations, as suggested by Arellano and Bond (1991) and Blundell and Bond (1998), will be used. The dynamic panel data method also considers the potential problems of endogeneity and reverse causality of the independent variables.

Baltagi (2005) presents certain advantages of using a panel data, namely, the possibility of controlling individual heterogeneity, greater informative power of data, greater variability, less collinearity between variables, greater freedom and more efficiency, better analysis of adjustment dynamics, the possibility of identifying and measuring effects which time-series or pure cross-section data do not capture.

Many economic relations are dynamic by nature and the dynamic panel allows for a better understanding of the adjustment process of these relations. This specification is characterized by the presence of a dependent variable lag among the independent variables (Baltagi, 2005). The following expression is assumed for the dynamic model:

[8] At Global Economic Monitor Commodities.

$$Y_{it} = \delta Y_{it-1} + \sum_{j=1}^{k} \beta X_{it}^{j} + \mu_i + \eta_{it} \; | \delta | < 1; \; i = 1,2,...,N; \;\; t = 1,2,...,T \tag{4}$$

where Y_{it} is a dependent variable, in this case, state i's exports for year t. X_{it} is the current value vector or lag of the explanatory variables. μ_i is the unobservable specific effect of the states, while $\mu_i \sim (0, \sigma_\mu^2)$ and the error term $\eta_i \sim (0, \sigma_\eta^2)$ are independent and identically distributed. The dynamic panel is estimated by the procedure known as the Generalized Method of Moments (GMM). The preference for estimator using the generalized method of moments is due to the fact that this method corrects the bias of fixed effects and also eliminates any endogeneity which could arise from the correlation of the specific effects of the states with the independent variables (BALTAGI *et al*., 2009). At the same time, it eliminates the problem of reverse causality in the estimation model.

Estimating equation (05) by the GMM-difference (GMM-dif) method, which eliminates the specific effects (μ_i), consists of the following specification:

$$Y_{it} - Y_{it-1} = \delta(Y_{it-1} - Y_{it-2}) + \beta(X_{it} - X_{it-1}) + (\eta_{it} - \eta_{it-1}) \tag{5}$$

In this model, the following moment condition is necessary, if there is to be orthogonality:

$$E[Y_{it-2}.(\eta_{it} - \eta_{it-1}] = 0 \quad \forall \quad t = 3,...,T \tag{6}$$

$$E[X_{it-2}.(\eta_{it} - \eta_{it-1}] = 0 \quad \forall \quad t = 3,...,T \tag{7}$$

In this case, Y_{t-2} is a valid tool in the first-order difference equation, as it is strongly correlated with $(Y_{it-1} - Y_{it-2})$ and is not correlated with the errors $(\eta_{it} - \eta_{it-1})$. As the model could be over-identified, the Sargan test should be applied to check the validity of the instruments chosen. In terms of the η_{it} errors, the GMM-dif estimate produces first-order correlated errors. Arellano and Bond (1991) present a hypothesis test that there is no second-order serial correlation in the first-difference equation disturbances. In this test the null hypothesis of the correlation of first-order errors (AR1) is not rejected, but the correlation of higher-order is.

Blundell and Bond (1998) claim that the lagged level of the series generates weak instruments for first-order difference estimates, especially when δ approaches a unit or when the specific effect variance increases, thereby expanding $\sigma^2_\mu/\sigma^2_\eta$. Using a study by Arellano and Bover (1995), they present the suggestion of estimating a system of equations using the GMM system (GMM-sys). This system uses both the first-order difference equation, already mentioned, and the equation in level with the first differences of the variables as a potential tool for this equation. The second part of the GMM-sys (regression level) depends on the following moment conditions:

$$E[(Y_{it-1} - Y_{it-2}).(\eta_{it})] = 0 \quad \forall \quad T = 3,...,T \tag{8}$$

$$E[(X_{it-1} - X_{it-2}).(\eta_{it})] = 0 \quad \forall \quad T = 3,...,T \tag{9}$$

Having made this brief presentation of the econometric procedure to be used, the proposed empirical model is now presented.

4.3. Empirical model

In order to investigate, in particular, the effects of Chinese income and real exchange rate on Brazilian states' agribusiness exports, using the Cronovich and Gazel (1998) model, we propose the following empirical specification for estimation.

$$\ln X_{i,j,t} = \beta_1 \ln wy^*_{i,t} + \beta_2 \ln w\varepsilon_{i,t} + \beta_3 \ln y_{i,t} + \beta_1 \ln pc_{wt} + u_{it} \tag{10}$$

Given that $u_{it} = \mu_i + \eta_{it}$ in which μ_i is the state's unobservable specific effect (fixed effect) and η_{it} the error term which represents economic shocks.

Finally, using the empirical model represented by equation (10), the results and analysis are presented.

5. Estimates and analysis

The results of the empirical model estimates presented (equation 10) through a basic pooled regression and static panel data are shown in Table 1. In estimates (1-4) the dependent variable is the volume of states' agribusiness exports to China and the regressions (3-4) the fixed effects are controlled. Firstly, the estimates (1-2) present baseline model in order to have a reference. Then, we estimate the regressions (3-4), in this model, the fixed effects capture state-specific factors, supply-side factors such as each state's abundance of natural resources (geography, arable land etc.). A positive sign is expected for all coefficients, for example, a high *wε* implies that exports are elastic in terms of the real exchange rate and depreciation in the exchange rate could lead to an increase in exports.

The estimated coefficients for the Chinese trade-weighted GDP (*wy**) presents a positive sign and are statistically positive, while the trade-weighted real exchange rate (*wε*), even with the expected sign, was not statistically significant in either estimate. In other words, from the estimates of the fixed effects model it can be seen that the real exchange rate is not a relevant variable for explaining the evolution of agribusiness exports to China in the period under consideration.

The estimated coefficients for the states' GDP (y) present the expected sign; however, in the third regression which includes the international agricultural commodity price variable (cp), it is no longer statistically significant at conventional levels. The estimated parameter for the international agricultural commodity price (cp) is statistically significant in the second regression only. Early evidence suggests that agribusiness exports from Brazilian states to China are determined by the Chinese GDP (demand side) and by the size of the states' GDP (supply side/production), and are not influenced by prices (*wε* and cp).

The main lesson to be taken from the results shown in Table 1 for Brazilian states' agribusiness exports to China is initial evidence of the importance of foreign income (Chinese GDP) on the demand side and the size of the states' economies on the production side.

Table 1 - Static model estimates of the fixed effects. Dependent variable: agribusiness exports (x)

Variables	MQO–Pooled		Panel	
	(1)	(2)	(3)	(4)
wy^*	1.300^{***}	1.323^{***}	1.010^{***}	1.011^{***}
	(0.089)	(0.085)	(0.052)	(0.053)
*w*ε	1.331	0.804	0.336	0.230
	(4.852)	(4.744)	(2.252)	(2.330)
y	1.180^{***}	1.192^{***}	0.708^{**}	0.497
	(0.066)	(0.065)	(0.280)	(0.383)
pc		-0.989^{**}		0.170
		(0.487)		(0.268)
Prob. test F			0.000	0.000
Hausman (p-value)			0.049	0.000
R^2	0,85	0.85		
R^2-within			0.927	0.927
Units			23	23
Observations	207	207	207	207

Source: Authors' calculations based on research data.
Notes: i) robust errors in brackets; ii) *** and ** indicate significance at 1 and 5%, respectively; iii) all estimates include one constant; iv) all variables are in Ln; regression (1 and 3) replicates the same specification as Cronovich and Gazel (1998).

To obtain more accurate results, new estimates are made considering the dynamic model; the crucial difference in relation to the fixed effect model is the inclusion of the lagged dependent variable (x_{t-1}) among the explanatory variables. Regressions 1-5 in Table 2 were estimated through a dynamic panel data procedure. In addition, variations were made in the basic specification to circumvent potential problems of endogeneity and reverse causality.

The estimate presented by regression (1) in Table 2 replicates the base model in its dynamic form and the results corroborate those found by the static model in which foreign income and the size of the states' economies are crucial for the dynamics of exports. The parameter of the real exchange rate variable (*w*ε) is not statistically significant and, it should be noted that the lagged export coefficient (x_{t-1}) is statistically significant which suggests that states' agribusiness exports could possibly have a dynamic component – a past history that matters.

Regressions 2-3 were carried out considering the lags in the real exchange rate and states' GDP states (see Cronovich and Gazel, 1998). The results show that the export lag is not statistically significant, while foreign income continues to have the expected sign and is statistically significant, irrespective of specification. The parameter of the states' GDP variable continues to have the expected sign and is significant both at variable level as in lag (regression 3).

The results presented by regressions 4 and 5 are obtained with the inclusion of the international agricultural commodity price (cp) variable in the specification (see Vieira and Haddad, 2011). It can be seen that the export lag (x_{t-1}) presents the expected sign and is statistically significant in model 4 while in model 5 the sign is changed and is no longer significant. The foreign income parameters (y*) corroborate the previous results in the same way as the states' GDP. The variables representing prices - real exchange rate (*w*ε) and international commodity prices (cp) – are not statistically significant, irrespective of specification. Results for the real exchange rate are consistent with those presented by Stockly *et al.* (2011) and Mortatti *et al.* (2011).

Table 2 - Dynamic model estimates - GMM. Dependent variable: agribusiness exports (x)

	Panel				
Variables	(1)	(2)	(3)	(4)	(5)
$x_{(t-1)}$	0.179*	-0.00585	-0.0727	0.179*	-0.0727
	(0.0949)	(0.371)	(0.370)	(0.0949)	(0.370)
wy^*	1.257***	1.231***	1.281***	1.257***	1.281***
	(0.203)	(0.217)	(0.217)	(0.203)	(0.217)
$w\varepsilon$	-4.930			-4.930	
	(8.626)			(8.626)	
$w\varepsilon_{(t-1)}$		16.48	23.90		23.90
		(37.88)	(37.34)		(37.34)
y	0.988***	1.245***		0.988***	
	(0.301)	(0.446)		(0.301)	
$y_{(t-1)}$			0.998**		0.998**
			(0.431)		(0.431)
pc				1.001	
				(1.143)	
$pc_{(t-1)}$					2.406
					(4.106)
dummy time	yes	yes	yes	yes	yes
AR(2)	0.175	0.154	0.110	0.175	0.110
Validity of instruments					
Sargan	0.760	0.875	0.930	0.760	0.930
Hansen-Diff	0.873	0.941	0.962	0.873	0.962
Number of instruments	13	13	13	13	13
Units	23	23	23	23	23
Observations	184	184	184	184	184

Source: Authors' calculations based on research data.
Notes: i) robust errors in brackets; ii) ***, ** and * indicate significance at 1, 5 and 10%, respectively; iii) all estimates include one constant; iv) all variables are in Ln; v) collapsing estimates to control the number of instruments.

The results in Table 2 (GMM System) are conditioned to the second-order autocorrelation test AR (2), and the validity of the instruments. The values (*P*-value) of the AR (2) test confirm that second-order autocorrelation is not a problem in the estimated models. And according to the statistics of both the Sargan and Hensen-diff tests it can be seen that the instruments are validated.

The dynamic estimates showed that the coefficients of foreign income (Chinese income) and the states' GDP recorded a greater economic impact (in terms of magnitude) when compared with the static estimates (Table 1). This result indicates that the static model could underestimate the economic impact of evolution in their respective variables. With the exception of the lagged exports (x_{t-1}), the coefficients showed stability in both sign and statistical significance, irrespective of specification which shows that the results are consistent.

In summary, the results of this research can be divided into two points: firstly, Brazilian states' agribusiness exports to China are not influenced by price (real exchange rate or international agricultural commodity price); and secondly, exports respond elastically to the growth of Chinese income and positively and inelastically to the states' supply capacity. This can be explained because in the 2000s Chinese growth was very impressive and exerted great pressure on the demand for agribusiness products even in a situation of increased prices, thus the sector's exports grew in accordance with the states' capacity.

6. Conclusions

This research set out to empirically analyze the determinants of Brazilian agribusiness exports to the Chinese market. To do so, estimates were made using the static and dynamic data panel based on agribusiness export data from Brazilian states.

The results are consistent with the literature and suggest that agribusiness exports from Brazilian states are influenced both by foreign income (Chinese) on the demand side, or by the states' GDP on the production/supply side. It can be seen that exports are elastic in relation to income and inelastic in relation to the states' GDP. In addition, it was found that both the real exchange rate and the international agricultural commodity price were not relevant in determining agribusiness exports in the period under consideration.

Although when the estimates were made using the static panel, the parameters were seen to be sensitive to change in specification, but when the dynamic procedure was considered, parameters remained stable. This confirms the consistency of the coefficients estimated for foreign income and states' output. With regard to the dynamic component, export lag, the coefficient was not stable for the different specifications, therefore not allowing for reliable considerations in relation to it.

In the Brazilian literature, the debate seeks to understand the determinants of agribusiness exports to China; however, empirical studies use the time-series method. This does not take the heterogeneity of the Brazilian states into account and could thus lead to results which over or underestimate the economic impacts on exports. Thus this study could contribute in two ways to the discussion on the issue: firstly, by constructing a state-specific, export-weighted variable; and secondly, estimating through dynamic panel data pointing to foreign income and the states' GDP as key determinants of agribusiness exports. Finally, in terms of promotion policies for agribusiness exports the results suggest that, as well as exchange influence, other policies should be drawn up.

Further research could be carried out including EU countries, USA, Japan and the Middle East, vital traders on the international market and trading partners of Brazilian agribusiness. Such research could lead to a greater understanding of the link between Brazilian agribusiness and China.

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ANNEX A

Table 3 – Agribusiness exports (in US$) of the selected states to China and the world market. Period: 2003 and 2011.

States	2003	2011	2003	2011
	Destination China		Destination World market	
AC	1016425	293924	4635584	9375128
AM	1111931	116281	17904115	19113846
BA	58554487	1225827580	722832114	3730089487
CE	107472	41937275	434551745	694222089
DF	179832	24469609	8901591	104338477
ES	92299024	321729354	941898041	1862187223
GO	49002737	1021217445	802760774	3000506859
MA	17048819	286923735	144248593	623811823
MT	282415173	3511348882	1789129930	7369814602
MS	29315364	771111514	336924173	2071492573
MG	114395668	597216307	1680894653	7687622720
PA	51850045	106669746	548919627	1108111001
PR	625095012	3169733402	3874287872	8458483028
PE	160724	4524651	220968542	547170990
PI	335903	60951043	51554644	143336688
RJ	1063838	8889	68755962	146267746
RN	55435	177462	212720424	184523637
RS	690791717	3193590095	4002610219	7790305513
RO	23411433	5546682	69600805	221350082
RR	225416	16077	1647617	7727796
SC	25120007	338189010	1861361715	3674191831
SP	171330236	1700661482	4588873540	12264301197
TO	8931865	107022984	42017539	414257172
Total	**2243818563**	**16489283429**	**22427999819**	**62132601508**

Source: *Agrostat* – MAPA (2014).

Note: i) states excluded from the sample: Alagoas, Amapá, Paraíba and Sergipe.