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The interplay of export supply and the real exchange rate. Evidence for Mercosur exports to the EU.

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Abstract

This paper applies a dynamic macroeconomic trade model to assess Mercosur-European Union trade. Looking at export supply of Mercosur countries (the four formal members plus Chile), the role of the real exchange rate, income and the income-absorption surplus or deficit are evaluated. Special emphasis is put on the reaction of exports with respect to changes of the real exchange rate. The model is tested for a sample of five countries (Argentina, Brazil, Chile, Paraguay and Uruguay) over the period of 1961-1996. A panel data analysis is used to disentangle the time invariant country-specific effects and to capture the relationships between the relevant variables over time. We find that the fixed effect model is to be preferred to the common effect model. The variables income and income-absorption surplus are found to be important determinants of trade flows. The real exchange rate has a positive and significant impact on export supply in the long-term, whereas current and past changes in the real exchange rate seem to play no role for current total export trade in the short-and medium-term. Having this latter time horizon, it could be shown that Mercosur's total exports react extremely parsimoniously and slowly with respect to changes in the real exchange rate. This phenomenon could be due to the large share of agricultural and forestry products in Mercosur's exports.

Key words: export supply; exchange rates; dynamic panel analysis

JEL classification: F14

1. Introduction

A very recent example of North-South integration is the EU-Mercosur trade agreement. The first negotiations started in 1995 with the signing of an Interregional Framework Agreement aimed to foster economic co-operation and closer trade relations between the two regional blocks. A further objective was the creation of a FTA in the year 2005. On the side of the EU, incentices to engage in substantive negotiations with Mercosur will depend closely on the consolidation and progress recorded by the Mercosur as a customs union. On the side of Mercosur, access to the EU market and the attraction of foreign direct investment are incentives playing a major role to engage into FTA negotiations with the EU.

Since its creation Mercosur has faced an extremely demanding agenda of extra-regional trade negotiations. It is considered an emerging market offering good investment opportunities, with a population over two hundred millions inhabitants (it represents half of the population of Latin America and Caribbean together) and an extension of almost 12 million squared kilometers. Mercosur has probably more to gain by joining the EU in a FTA rather than negotiating with North America, since

Mercosur member countries already have relatively free access to the North American market. This papers intends to evaluate the export potential of the Mercosur and of Chile as an associated country by examining the determinants of exports from those countries to the European Union (EU) in the period of 1961-1996. Especial emphasis will be put on the interplay between Mercosur's export supply and the development of the real exchange rate, i. e. of relative prices. In our work the price elasticity of export supply will be analyzed for the long, medium and short-run. Since a FTA between the EU and the Mercosur implies, among other things, a lowering and abolishing of tariffs, more competitive real exchange rates (increased exchange rates) are to be expected for the two blocks. Furthermore, the interplay between export supply and business cycle variables, such as production capacity and the income-absorption surplus (or deficit) will be subject to analysis. Both aspects, real exchange rate and business cycle, are then integrated into our trade model which is to be estimated by methods related to the pooled analysis. The long-term model has a very simple structure, serving as a benchmark where lagged reactions do not exist. At the center of interest is the short-to medium-term model. This model assumes that the supply of exports in each of the five Mercosur countries adjusts with lags to changes in the real exchange rate. In our analysis, the lag structure is depicted by a polynomial.

There are several novelties in our approach. First, this is the first attempt that does not run the regressions in the usual first difference form, but in a 'soft' first difference form which leaves more information in the series. Second, to our knowledge another novelty in our pooled analysis consists in producing a dynamic model that uses a lag structure justified by the data. Finally, the trade relations between Mercosur and the EU have not yet been deeply analyzed. Only a few attempts have been made in this direction in Martínez-Zarzoso and Nowak-Lehmann D. (2001) and Bulmer-Thomas (2000).

Dynamic modeling enables us to judge whether - by looking at the export side - an adjustment with lags (i.e. a type of exchange rate hysterisis) is characteristic for total exports in the short-and medium run. If this is so, appreciations of the real exchange rate are not quite as harmful for the trade balance and depreciations (due to tariff reductions) will not be quite as effective under this time horizon. However, one should not overlook the fact that hysterisis of total exports might be due to the large share hold by the exchange rate-inelastic agricultural and forestry products. In other words, the real exchange rate could still improve - via a real depreciation - or impede - via a real appreciation - the international competitiveness of price-elastic goods, as e.g. manufactured exports.

The remainder of the paper is organized as follows. In Chap. 2 we discuss the theoretical model. In Chap. 3 we derive the empirical equations for estimation purposes. Chap. 4 provides the estimation results and finally Chap.5 concludes.

2. Modeling the (lagged) relationship between exports and the real exchange rate

Economic theory does not cease to emphasize the role of relative prices for the production of exportables. An early model that reflects this relationship is the Australian model such as propagated by Salter (1959) and Swan (1960)[1]. This model has been refined and augmented by several economists[2]. Goldstein and Khan (1978) added the variable domestic production capacity, thus generating eq. (1).

(1) $\log X^s_t = \beta_0 + \beta_1 \log(PX/P)_t + \beta_2 Y^*_t$

where

X^s = quantity of exports supplied

PX = price of exports

P = domestic price index

Y* = logarithm of an index of domestic capacity

Other authors emphasize the role played by capacity utilization and domestic demand, variables which try to capture the correlation between strong export growth and the presence of large unemployment of domestic resources (Faini, 1994).

We follow this line of thought by incorporating the income-absorption surplus/deficit as an additional determinant of the supply of exports.

The long-term model in which exports, production capacity, income-absorption surplus/deficit and the real exchange rate should stay in line with each other, i.e.,

[1] See Dornbusch (1980).

[2] Compare Beenstock et al., 1994; Ceglowski, 1997; Faini, 1994; De Gregorio, 1984; Khan and Knight, 1988; Lukonga, 1994; Moreno, 1997; Newman, 1995; Rodgers, 1998; Wang, 1998:

they should be co-integrated, is then formulated as co-integrating regression (2) in logs[3].

(2) $lx_{it} = \alpha_i + \beta\ lyx_{it} + \gamma\ tb_{it} + \delta\ ler_{it} + \varepsilon_{it}$

where:

i stands for Mercosur country i

t stands for year, t = 1961-1996

α_i denotes individual effects

lx_{it} = supply of exports in real term in logs from country i in period t

lyx_{it} = income of the exporting country in real terms and in logs; it serves as an indicator of the production capacity of the exporting economy

tb_{it} = income-absorption surplus (positive value) or deficit (negative value) of the exporting country in percentage

ler_{it} = index of the real exchange rate with 1995 = 100; an increase implies a devaluation

The expected sign for α_i can be positive or negative, whereas the expected signs for β, γ, δ are all positive. An increase in the production capacity of the exporting country is seen to translate into a reinforced production of export goods. A country that possesses an income-absorption surplus will get rid of the surplus by exporting and a real exchange rate depreciation will result in an increased supply of exports due to strengthened price competition.

One should note that the use of this type of static models is considered adequate when the long-term relationship between exports and the real exchange rate is under scrutiny (Pindyck and Rubinfeld, 1991).

However, static modeling ceases to be useful when the short- to medium term relationship between the supply of exports and the real exchange rate is at the center of interest. Especially, when we have reasons to assume that exports react and respond with lags to changes in the real exchange rate. In this case dynamic modeling is required.

The augmented Australian model can be made dynamic by allowing for lagged relationships between the dependent variable (export supply) and the independent

[3] The original form of the model is multiplicative. By taking logs the model is linearized and made estimable. Since the variable tb_{it} can also take on negative values, the log would not be defined and therefore the variable remains unaltered.

variable under special scrutiny, in our case the real exchange rate. A distributed lag model can be built (Nowak-Lehmann D., 1997)

(3) $lx_{it} = \alpha_i + \beta \; lyx_{it} + \gamma \; tb_{it} + \sum_{k=0}^{K} \delta_k \; ler_{it-k} + \varepsilon_{it}$

with:

$\sum_{k=0}^{K} \delta_k \; ler_{it-k} = \delta_0 ler_{it-0} + \delta_1 ler_{it-1} + ... + \delta_K ler_{it-K}$

where:

k denotes the length of the lag in years and δ_k stands for the coefficient belonging to the real exchange rate lagged by k periods

The response of exports with respect to changes in the real exchange rate can take on a multitude of different shapes depending on δ_k. The selection of the 'right' shape has to be derived from the data. An overview of the most common lag structures is offered by Nowak-Lehmann D. (1997). A popular lag structure is the geometric lag which is characterized by a given form (Kelejian and Oates, 1989; Greene, 2000). The gamma lag model is quite unknown, also of a given form, and non-standard to estimate[4] (Schmidt, 1974). Only if the data follow this form, the use of the geometric or the gamma lag is justified. The transfer function model, which allows to model any lag structure suggested by the data, is much more flexible. In this case the lag structure is described by a polynomial in the numerator and a polynomial in the denominator (Box and Jenkins, 1976; Greene, 2000). However, most standard econometric software does not support the estimation of the transfer function. Therefore, the 'simple' polynomial (polynomial only in the numerator) was chosen (Greene, 2000). It is usually the most convincing method of modeling lags. By determining the order of the polynomial and the length of the lag, one can basically model any lag structure suggested by the cross-correlations between the impulse variable (in our case, the real exchange rate) and the response variable (in our case, export supply).

[4] It has to be estimated either with the Maximum-Likelihood method or a non-linear least square procedure.

3. Estimating the dynamic adjustment process

Dealing with macroeconomic panel data with a long time dimension time (T=36), one must be aware of variables with undesirable time series properties. One such property is the non-stationarity of the series. The notion of non-stationarity entered the econometric literature about 20 years ago. Earlier the time series properties of the series were - in general - not yet considered critical when running regressions.[5] Once the spurious regressions effect of non-stationary series had been discovered, the formulation of the regression equations in first difference form became the method of choice.

In a first step therefore, we ran tests on non- stationarity for each of the five countries Argentina (AR), Brazil (BR), Chile (CH), Paraguay (PA) and Uruguay (UR) separately. As shown in Table 1, all the variables in logs, namely lx, lyx, tb and ler turned out to be non-stationary, but cointegrated.[6] Since transformation of the series into the usual first difference form[7] has the disadvantage of wasting long-run information, a different approach was followed (Greene, 2000). In a second step, we freed the variables from their time trend and ran the regression in a 'soft' first difference form. This is achieved by applying the FGLS-method (Feasible General Least Squares-method). This procedure works as follows. The coefficient of autocorrelation (ρ) between the error terms ε_{it} and ε_{it-1}, reflecting non-stationarity of the series lx_{it}, lyx_{it}, tb_{it} and ler_{it}, is computed in the original model (2) , leading to $\hat{\rho}$.

(4) $\varepsilon_{it} = \rho\varepsilon_{it-1} + v_{it}$

Then 'soft first differences' with $\hat{\rho}$ for all the series are generated. The new series carry the supplementary z. We get:

$lxz_{it} = lx_{it} - \hat{\rho}\, lx_{it-1}$

$lyxz_{it} = lyx_{it} - \hat{\rho}\, lyx_{it-1}$

$tbz_{it} = tb_{it} - \hat{\rho}\, tb_{it-1}$

$lerz_{it} = ler_{it} - \hat{\rho}\, ler_{it-1}$

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[5] One must mention the famous work of Granger and Newbold (1974) who discovered spurious regressions to result when running regressions with non-stationary series. However, this information entered the econometric textbooks only in the middle of the 80s or at the beginning of the 90s.

[6] Cointegration implies that the supply of exports and its macroeconomic determinants are in long-run equilibrium. However, when analyzing short- to medium term economic behavior, this information is of little help.

[7] In the usual first difference form the relationship: $y_t = \rho\ y_{t-1} + u_t$ is characterized by ρ =1. In the soft first difference form ρ is determined by the autocorrelation between the error terms. It will usually be much smaller than 1 (in our case ρ =0.55)

As far as the issue of dynamic modeling is concerned, an early presentation of dynamic models using panel data can be found in Anderson and Hsiao (1981 and 1982). In the existing literature on panel data dynamic modeling is achieved by incorporating a lagged endogenous variable. This lag form is named geometric lag or Koyck lag. It is very simple from a superficial perspective, but is burdened with estimation problems. Estimation problems occur when the error terms loose their desirable properties. Autocorrelation of the disturbance terms is one of the problems to be dealt with (Kelejian and Oates, 1989). A documentation of those estimation problems is given in the article of Hansen (1999). Arellano and Bond (1991) estimated the regression equations in first difference form via the Generalized Instrumental Variable Estimator (GIVE) or the General Method of Moments (GMM). Hansen (1999) showed by means of Monte-Carlo experiments that these estimators are more biased and less reliable as a 'bias-corrected Least-Squares-Dummy-Variable (LSDV)-estimator which is to be considered 'best' in the class of problematic estimators.

In this paper we use a different approach, using an estimation method that avoids the problems mentioned above (Nowak-Lehmann D., 1997). First of all, the geometric lag is presumptive in form and can only to be justified if the underlying sample supports this shape. Second, the geometric lag can also be represented by a polynomial. There is a theorem in mathematics that states that, under general conditions, a curve may be approximated by a polynomial (see Kelejian and Oates, 1989). This theorem is used to determine the lag structure in eq. (5).

(5) $lxz_{it} = \alpha_i(1-\hat{\rho}) + \beta\ lyxz_{it} + \gamma\ tbz_{it} + \sum_{k=0}^{K}\delta_k lerz_{it} + v_{it}$

From a practical point of view, the best way to determine the lag structure between the response variable (lxz_t) and the impact variable ($lerz_t$) is by looking at the cross correlations between lxz_t and $lerz_t$ in each of the Mercosur countries.

Our single equation cross correlations suggest a maximum lag length of three years. As far as the shape is concerned, it can be depicted best by a polynomial of degree 1. Taking this approach we get:

(6) $\delta_k = d_0 + d_1 k$ with k= 0, 1,....K

Equation (6) helps to derive indirect estimates for the δ_k

(6a) $\delta_0 = d_0$

(6b) $\delta_1 = d_0 + d_1$

(6c) $\delta_2 = d_0 + 2d_1$

(6d) $\delta_3 = d_0 + 3d_1$

Once the lag structure has been modeled, this information has to enter equation (5), yielding equation (7).

(7) $lx_{it} = \alpha_i(1-\hat{\rho}) + \beta\ lyxz_{it} + \gamma\ tbz_{it} + d_0\ (\sum_{k=0}^{K} lerz_{it-k}) + d_1\ (\sum_{k=0}^{K} k * lerz_{it-k}) + v_{it}$

where:

$\alpha_i(1-\hat{\rho})$ becomes a_i

$\sum_{k=0}^{K} lerz_{it-k}$ becomes $z1lerz_{it}$

$\sum_{k=0}^{K} k * lerz_{it-k}$ becomes $z2lerz_{it}$

The model thus simplifies to

(8) $lx_{it} = a_i + \beta\ lyxz_{it} + \gamma\ tbz_{it} + d_0\ z1lerz_{it} + d_1\ z2lerz_{it} + v_{it}$

Equation (8) can be estimated by the techniques available in the pooled analysis. The parameters δ_0, δ_1, δ_2 and δ_3 are computed according to formulas (6a)-(6d).

4. Empirical evidence

Since the number of cross sections in our study was small, we had the opportunity to do a pre-study by running single regressions for each of the five Mercosur countries and to use the information from our five different time series analyses.

By looking at each single country we could also detect and determine the shape of the lagged relationship between the supply of real exports and the real exchange rate. [8] This information was crucial for selecting a polynomial that was able to reflect the lag structure suggested by the data.

The long-run model (eq. (2)) was estimated in 'soft difference form', i. e. with stationary series carrying the suffix 'z' [9]. In general this sort of caution is not necessary as long as the non-stationary variables are cointegrated in the long-run.[10] Our original variables proved to be non-stationary, but cointegrated, i.e. in long-run

[8] The series have to be stationary for this purpose. We used the variables in 'soft first difference form'.
[9] The series with the suffix 'z' were all stationary, except for lyxz.
[10] In the long-run, it is considered viable to run regressions with non-stationary series as long as they are cointegrated.

equilibrium (see Table 1). Nonetheless, we decided to estimate the long-run equation with our 'z'- series since the period of 1961-1996 might not be long enough to classify for the long-run in economic terms. Even though the R^2 calculated in our approach will be smaller than in the regression runs with non-stationary series, it provides a more honest measure of goodness of fit.[11]

The short-to medium run equation was also run with stationary series which are a 'must' in this case. The new series with the suffix 'z' were all stationary, except for lyxz, according to the Phillips-Perron test. (see Table 2).

In the pooled analysis framework with 174 unbalanced observations, 5 cross sections and an adjusted sample running from 1962-1996, we then applied the Hausman test to check for endogenity of the regressors. The Hausman test did not reveal any problem of endogenity of our 'right hand side' variables: lyxz, tbz, and lerz. Results are shown in Table 3.

Finally we could start to estimate the long-run equation (2). Different specifications concerning the constant term (common effect, fixed effects, and random effects) were tested against each other (see Table 4).

The fixed effects model proved to be superior to the common effect model. Furthermore, the random effects model had to be ruled out since the number of cross sections has to be bigger than the number of coefficients to be estimated.

The fixed effects model had a clear advantage over the common effect model which was rejected. Within the fixed effects model the results of the GLS (General Least Squares with cross section weights)[12] and of SUR (Seemingly Unrelated Regression)[13] were very similar (see Table 5 and 6). (Adjusted) R^2 was 0.68 (0.66) under GLS and SUR. AR(1) terms, not being significant, were not plugged into the model.

The results show that the domestic production capacity (lyxz), the income-absorption surplus (tbz) and the real depreciations (lerz) do all have a positive and significant (as expected) impact on export supply in the long-run. It should be pointed out that the real exchange rate (lerz) is crucial for export supply, as suggested by neoclassical theory. This result differs from the estimations in the short-and medium term, as will become evident in the next section.

[11] The R^2 in the GLS and SUR estimation, based on non-stationary, but cointegrated series, was 0.82. The R^2 in the GLS and SUR estimation, based on stationary series, was 0.68.

[12] GLS gives different weights to the cross sections (5 Mercosur countries) and might therefore be more meaningful than unweighted Pooled Least Squares estimation.

[13] SUR takes autocorrelation between the error terms of the cross sections (countries) into account.

In contrast to the long-run model, a polynomial lag (of degree 1 and the maximum length of 3 years) was built into the short-and medium-term model (based on eq. (3)). The model was estimated using stationary series with the suffix z and eq. (8). This required some variable transformations.

In the short-and medium-term scenario, the GLS and the SUR estimation led also to very similar results as displayed in Table 7 and 8. The domestic production capacity (lyxz) and the income-absorption surplus (tbz) had a positive and significant impact on export supply (lxz). The incorporation of the adjustment lag between export supply (lxz) and the real exchange rate (lerz) improved clearly the explanatory power of the model. (Adjusted) R^2 increased (from 0.68 (0.66 in the perfect adjustment version)) to 0.92 (0.91) under both GLS and SUR. Autocorrelation was corrected in both estimations via insertion of an AR(1)-term, which proved to be significant.

Table 9 shows that current and past changes in the real exchange rate did neither in the GLS nor in the SUR estimation have a significant impact on exports. Relying on the formulas (6a)-(6d) we obtained the following results: $\delta_0 = 0.006, \delta_1 = 0.07, \delta_2 = 0.13, \delta_3 = 0.19$ in GLS with increasing significance of the coefficients (even though overall still insignificant) and in SUR: $\delta_0 = -0.02, \delta_1 = 0.05, \delta_2 = 0.12, \delta_3 = 0.19$, again with increasing significance (even though still insignificant).

To sum up, the results of the previous sections show that the exchange rate has a positive and significant impact in the long-run. In the short-to medium term, in contrast, no such impact can be found. Even though the coefficients have the right sign (with the exception of δ_0 in SUR), they are not significant. However, one should note a slightly increasing exchange rate elasticity from δ_0 to δ_3. I. e. Exchange rate changes that occurred three periods back have a bigger impact on exports than exchange rate changes that took place two, one or zero periods back. One important reason for the sluggish reaction of exports vis-à-vis the exchange rate is the large share of agricultural, forestry and fishery goods in Mercosur's exports to the EU. According to OECD data their share was 36.78 % in 1996.[14] The process of producing these products might take several years, especially investment to enlarge the production of e.g. wood, fruit, salmon, beef (Nowak, 1989). Besides, investigations on manufactured exports clearly show a positive and significant impact

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[14] This figure refers to the sectors 00 to 09 according to OECD classification.

of current and past exchange rates on exports (Nowak-Lehmann D., 1997). Manufactured exports are in general considered exchange rate elastic. A comprehensive study of Mercosur's exports and its sub-groups should follow in the future.

5. Conclusions

The study gives some insights into the role played by the real exchange rate, the production capacity and the income-absorption surplus, in explaining total Mercosur's exports to the EU. In the long-run all these factors have a positive and significant impact on total exports and eventually total export growth. In the short- and medium run, production capacity and income-absorption surplus keep their relevance, whereas the real exchange rate ceases to be significant. To be more precise, current and past developments of the real exchange rate discontinue to have a significant impact on exports (total exports). We attribute this outcome to the high share of the 00-09 categories in Mercosur's exports to the EU that must be considered exchange rate inelastic.

The explanatory power of the model has been improved by building-in reaction lags. Certainly more dynamic econometric studies in the trade area are needed to evaluate and specify our results.

Evidence from older studies revealed the exchange rate elasticity of manufactured goods respectively exports. More investigations, especially on a sector-level, are needed to develop a strategy to improve international competitiveness of the exporting sectors.

Our work shows that the real exchange continues to be an important determinant of international competitiveness and attention should be paid to the development of the real exchange rate. This recommendation should be followed whenever the promotion of manufactured exports is considered important, which is the case in the Mercosur countries and Chile. On an international level the negotiations on a FTA between the Mercosur and the EU should be given a high priority.

Table 1: Test results of the original series in logs, country by country

Argentina

Test of stationarity with the Phillips-Perron-unit root test		
lx_arue	non-stationary ($\alpha = 1\%$)	PP[*] = -2.87
ler_arue	non-stationary ($\alpha = 1\%$)	PP = -3.55
lyx_arue	non-stationary ($\alpha = 1\%$)	PP = -1.84
tb_arue	non-stationary ($\alpha = 1\%$)	PP = - 3.12
Test of cointegration with the Johansen-cointegration test		
The Argentine series are cointegrated. The test indicates 1 cointegration equation at 5% significance level.		

Brazil

Test of stationarity with the Phillips-Perron-unit root test		
lx_brue	non-stationary ($\alpha = 1\%$)	PP = -4.21
ler_brue	non-stationary ($\alpha = 1\%$)	PP = -3.11
lyx_brue	non-stationary ($\alpha = 1\%$)	PP = -0.73
tb_brue	non-stationary ($\alpha = 1\%$)	PP = - 1.99
Test of cointegration with the Johansen-cointegration test		
The Brazilian series are cointegrated. The test indicates 2 cointegration equations at 5% significance level.		

[*] PP stands for Phillips-Perron test statistic

Chile

Test of stationarity with the Phillips-Perron-unit root test		
lx_chue ler_chue lyx_chue tb_chue	non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$)	PP* = -2.84 PP = -2.79 PP = -0.59 PP = - 2.84
Test of cointegration with the Johansen-cointegration test		
The Chilean series are <u>not</u> cointegrated. The test rejects any cointegration at 5% significance level.		

Paraguay

Test of stationarity with the Phillips-Perron-unit root test		
lx_paue ler_paue lyx_paue tb_paue	non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$)	PP = -1.98 PP = -2.19 PP = -1.14 PP = - 2.05
Test of cointegration with the Johansen-cointegration test		
The Paraguayan series are cointegrated. The test indicates 1 cointegrating equation at 5% significance level.		

Uruguay

Test of stationarity with the Phillips-Perron-unit root test		
lx_paue ler_paue lyx_paue tb_paue	stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) non-stationary ($\alpha = 1\%$) stationary ($\alpha = 1\%$)	PP = -4.49 PP = -3.28 PP = -2.28 PP = - 6.97
Test of cointegration with the Johansen-cointegration test		
In the Uruguayan sample the series are integrated of different orders. Some are I(0), i.e. stationary, some are I(1), i.e. non-stationary.		

* PP stands for Phillips-Perron test statistic

Table 2: Test results of the variables, in 'soft difference form', country by country

Argentina

Test of stationarity with the Phillips-Perron-unit root test		
lxz_arue	stationary ($\alpha = 1\%$)	PP* = -3.84
lerz_arue	stationary ($\alpha = 10\%$)	PP = -3.52
lyxz_arue	stationary ($\alpha = 1\%$)	PP = -4.41
tbz_arue	stationary ($\alpha = 1\%$)	PP = - 5.52

Brazil

Test of stationarity with the Phillips-Perron-unit root test		
lxz_brue	stationary ($\alpha = 1\%$)	PP = -8.18
lerz_brue	stationary ($\alpha = 1\%$)	PP = -4.20
lyxz_brue	non-stationary ($\alpha = 1\%$)	PP = -1.18
tbz_brue	stationary ($\alpha = 1\%$)	PP = - 4.42

Chile

Test of stationarity with the Phillips-Perron-unit root test		
lxz_chue	stationary ($\alpha = 1\%$)	PP = -4.28
lerz_chue	stationary ($\alpha = 1\%$)	PP = -4.51
lyxz_chue	non-stationary ($\alpha = 1\%$)	PP = -1.90
tbz_chue	stationary ($\alpha = 1\%$)	PP = - 4.66

* PP stands for Phillips-Perron test statistic

Paraguay

Test of stationarity with the Phillips-Perron-unit root test		
lxz_paue	stationary ($\alpha = 1\%$)	PP = -3.94
lerz_paue	stationary ($\alpha = 5\%$)	PP = -3.24
lyxz_paue	non-stationary ($\alpha = 1\%$)	PP = -1.34
tbz_paue	stationary ($\alpha = 1\%$)	PP = - 3.72

Uruguay

Test of stationarity with the Phillips-Perron-unit root test		
lxz_paue	stationary ($\alpha = 1\%$)	PP = -8.45
lerz_paue	stationary ($\alpha = 1\%$)	PP = -6.16
lyxz_paue	non-stationary ($\alpha = 1\%$)	PP = -2.78
tbz_paue	stationary ($\alpha = 1\%$)	PP = - 14.54

Table 3: Results of the Hausman test

Variables tested	Endogenity?	Coefficient of 'auxiliary' residual	t-value
Production capacity	lyxz exogenous	-0.30	-1.17
Income-absorption surplus	tbz exogenous	0.03	0.53
Real exchange rate	lerz exogenous	-0.64	-1.66

Note:
EViews, the statistical software used, does not perform the Hausman test in the classical way by comparing the β-vector under TSLS and under OLS. EViews rather runs an auxiliary regression, a method proposed by Davidson and MacKinnon (1989, 1993). For this purpose the variable that is suspect of being endogenous is regressed on all the exogenous variables in the equation. Then the residual of this auxiliary regression is plugged into the original regression equation. If the coefficient of this residual is significantly different from zero, the variable is considered being endogenous. If the residual is not significant, the OLS estimates are consistent and the variable is taken for exogenous (EViews, Version 3, User's Guide, 360-362).

Table 4: Which model applies? Common effect, fixed effects or random effects model?

	Common effect model	Fixed effects model	Random effects model
Perfect adjustment model (long-run model)	H_0 : there is a common constant, i.e. the fixed effects are equal to each other $\Rightarrow$ common effect model has to be rejected in favor of the fixed effects model test results: - in GLS: F-statistic = 16. 75 (probability:0.00) - in SUR: F-statistic = 14.72 (probability: 0.00)		the random effects model has to be ruled out since the number of cross sections is smaller than the number of coefficients to be estimated
Imperfect adjustment model (short- to medium run model)	H_0 : there is a common constant, i.e. the fixed effects are equal to each other $\Rightarrow$ common effect model has to be rejected in favor of the fixed effects model test results: - in GLS: F-statistic = 39.50 (probability:0.00) - in SUR: F-statistic = 49.83 (probability: 0.00)		the random effects model has to be ruled out since the number of cross sections is smaller than the number of coefficients to be estimated

Table 5: Estimation results for the fixed effects perfect adjustment model using GLS

Right hand side variables	Variable names	Estimated coefficient	t-value
Production capacity	lyxz	1.03***	11.47
Income-absorption surplus	tbz	0.12***	4.78
Real exchange rate	lerz	0.22**	2.37
Fixed effects:			
Argentina	ARUE-C	-4.53***	-6.22
Brazil	BRUE-C	-4.68***	-6.19
Chile	CHUE-C	-3.98***	-6.09
Paraguay	PAUE_C	-3.94***	-6.71
Uruguay	URUE-C	-4.25***	-6.69
R^2 = 0.68 adjusted R^2 = 0.66 SSR = 31.22		Durbin-Watson stat. = 1.33 number of observations = 174 ***, **, * = significant at 1%, 5%, 10%	

Table 6: Estimation results for the fixed effects perfect adjustment model using SUR

Right hand side variables	Variable names	Estimated coefficient	t-value
Production capacity	lyxz	1.01***	10.43
Income-absorption surplus	tbz	0.13***	5.28
Real exchange rate	lerz	0.22**	2.42
Fixed effects:			
Argentina	ARUE-C	-4.32***	-5.43
Brazil	BRUE-C	-4.47***	-5.40
Chile	CHUE-C	-3.80***	-5.31
Paraguay	PAUE_C	-3.77***	-5.87
Uruguay	URUE-C	-4.07***	-5.87
R^2 = 0.68 adjusted R^2 = 0.66 SSR = 31.32		Durbin-Watson stat. = 1.35 number of observations = 174 ***, **, * = significant at 1%, 5%, 10%	

Table 7: Estimation results for the fixed effects polynomial lag model using GLS

Right hand side variables	Variable names	Estimated coefficient	t-value
Production capacity	lyxz	1.08***	12.91
Income-absorption surplus	tbz	9.85***	13.68
Transformed real ex-change rates[f]	z1lerz	0.01	0.12
	z2lerz	0.06**	2.17
Fixed effects:			
Argentina	ARUE-C	-5.31***	-8.31
Brazil	BRUE-C	-5.52***	-8.30
Chile	CHUE-C	-4.82***	-8.41
Paraguay	PAUE_C	-4.73***	-9.22
Uruguay	URUE-C	-5.10***	-9.46
R^2 = 0.92 adjusted R^2 = 0.91 SSR = 4.93		Durbin-Watson stat. = 1.91 number of observations = 154 ***, **, * = significant at 1%, 5%, 10%	

[f] of theoretical importance; see Table 9 for the information that is of practical importance

Table 8: Estimation results for the fixed effects polynomial lag model using SUR

Right hand side variables	Variable names	Estimated coefficient	t-value
Production capacity	lyxz	1.15***	15.11
Income-absorption surplus	tbz	10.07***	15.16
Transformed real ex-change rates[f]	z1lerz	-0.02	-0.36
	z2lerz	0.07***	2.64
Fixed effects:			
Argentina	ARUE-C	-5.79***	-10.56
Brazil	BRUE-C	-6.01***	-10.54
Chile	CHUE-C	-5.25***	-10.69
Paraguay	PAUE_C	-5.10***	-11.67
Uruguay	URUE-C	-5.49***	-12.00
R^2 = 0.92 adjusted R^2 = 0.91 SSR = 4.97		Durbin-Watson stat. = 1.89 number of observations = 154 ***, **, * = significant at 1%, 5%, 10%	

[f] of theoretical importance; see Table 9 for the information that is of practical importance

Table 9: Overview over the impact of the real exchange rate (RER), lagged and unlagged [f]

Impact of RER in the GLS estimation		
impact of RER of current year (unlagged)	$\delta_0 = 0.006$	t = 0.13
impact of RER of 1 year back (lagged by 1 year)	$\delta_1 = 0.07$	t = 0.88
impact of RER of 2 years back (lagged by 2 years)	$\delta_2 = 0.13$	t = 1.18
impact of RER of 3 years back (lagged by 3 years)	$\delta_3 = 0.19$	t = 1.36
Impact of RER in the SUR estimation		
impact of RER of current year (unlagged)	$\delta_0 = -0.02$	t = - 0.36
impact of RER of 1 year back (lagged by 1 year)	$\delta_1 = 0.05$	t = 0.55
impact of RER of 2 years back (lagged by 2 years)	$\delta_2 = 0.12$	t = 0.80
impact of RER of 3 years back (lagged by 3 years)	$\delta_3 = 0.19$	t = 0.86

Note: None of the coefficients of the lagged real exchange rates is significant. However, its significance is increasing with higher lag order (increasing t-values).

[f] These results are to be computed according to the formulas (6a)-(6d).

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Appendix

Data sources

CEPAL, Statistical Year Book for Latin America and the Caribbean. Various years. United Nation Publication:

-Bilateral trade Mercosur + Chile

OEA, America en Ciphers 1965, 1970:
-Bilateral trade Mercosur+Chile

WILKE, James, Statistical Abstract of Latin America, Vol. XVII University of California Los Angeles (1976):
-Bilateral trade Mercosur+Chile

BID, Intra-ALALC exports (grouped according to Standard International Trade Classification) Various years (1965-1969):
-Bilateral trade Mercosur+Chile

OCDE, International Trade by Commodities Statistics ITCS. CD ROM 1960-1996:
-Bilateral trade for MERC countries

World Bank, World Development Indicators CD ROM 2000:

-GDP

-GDP deflator.

-(Total exports and imports)/GDP

-Exchange rates against dollar

World Bank, World Data 1995 CD ROM:

-Germany data before 1990

Estimated data:
-Bilateral real exchange rate (base 1995)
-Exports deflator (base 1995)
-Exports in real terms (base 1995)
-Trade weight
-Germany data prior 1990