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JOINT MODELLING THE EFFECT OF
COMMODITY PRICES ON EXPORTS AND THE
REAL EXCHANGE RATE: THE CASE OF
ARGENTINA

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Joint modelling the effect of commodity prices on exports and the real exchange rate: the case of Argentina

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Abstract

Argentina's rapid economic growth during the last decade has been largely associated with the commodity export boom. However, growing exports may also have side effects from exchange rate appreciation. To take into account the direct and indirect commodity price effects on exports and exchange rate we estimate a system which also includes agricultural output and domestic consumption. We found two cointegration vectors which allow us to identify an export equation and a real factor equation for the exchange rate. Endogeneity of the variables are explicitly modelled. A simultaneous equilibrium correction representation is also estimated to study short-run dynamics.

JEL Codes: O13, C32

Keywords: natural resources, real exchange rate, commodity prices, system cointegration, equilibrium correction models, Dutch disease.

Resumen

El reciente crecimiento de Argentina se ha asociado en gran medida al incremento de las exportaciones de commodities. Sin embargo, crecientes exportaciones pueden tener efectos secundarios debido a apreciaciones del tipo de cambio. Para considerar los efectos de los precios de las commodities sobre exportaciones y tipo de cambio, estimamos un sistema que también incluye la producción agrícola y el consumo interno. Encontramos dos vectores de cointegración: una ecuación de exportaciones y otra del tipo de cambio real. Consideramos explícitamente la endogeneidad de las variables. Para estudiar la dinámica de corto plazo estimamos un modelo de corrección al equilibrio.

Códigos JEL: O13, C32

Palabras clave: recursos naturales, tipo de cambio real, precios de commodities, cointegración en sistemas, modelos de corrección al equilibrio, Enfermedad Holandesa

1 Introduction

The commodity boom over the last decade seems to open a new reality for developing economies that produce and export natural resources. For those facing the positive (and unusual long-lasting) shock on commodity prices, the resulting export growth of natural resources seems to be a blessing for their economies, but it has had also important “side effects” on the exchange rate, the economic structure and income distribution.

There exists a great debate about the relationship between commodity dependence and growth in developing countries. Some empirical studies show a negative impact of natural resources on economic growth (Sachs and Warner, 1995, 2001). This view has become generally accepted by those who believe that the negative link between commodity prices and economic growth is driven by the Dutch disease. Natural resources abundant countries generate large profits for their producers and have a real exchange rate appreciation raising the price of non-tradable goods relative to that of tradable goods, and thus increasing their returns to production relative to other tradable goods. Therefore, there are no incentives to invest in other tradable goods, resulting in a highly commodity-specialized economy.

Argentina is a historical well-known commodity producer and exporter; its exports still remain highly concentrated on a few raw materials and lightly processed primary products. By 2011, seven out of the ten top export items consist of raw materials such as soybean meal, soybeans, soybean oil, corn, wheat, gold and crude oil which account for 38% of total merchandise exports.¹

The rapid growth of the Argentine economy during the last decade has been largely associated with a commodity export boom. However, the economic expansion after 2001-2002 crisis was not only sorted by favourable international commodity prices, but by a more flexible exchange rate regime and an expansionary domestic aggregate demand (mainly private and public consumption). Since this change in domestic consumption should have affected the real exchange rate too, we carry out an empirical investigation of the factors explaining both exports and the exchange rate. The purpose of simultaneously modelling these key variables is to elucidate how commodity prices affect Argentine exports and real exchange rate. This approach is also useful to evaluate whether or not Argentina has been suffering the effects of a Dutch disease since the last decade.

Despite the fact that Argentina has a traditional reputation of being a resource-rich country largely based on agricultural production, the effect of commodity prices on the economy was not often analysed. As an exception, the work of Fanelli and Albrieu (2013) highlights the positive shock that Argentina experienced in the last decade on the terms of trade, although they warn about how fragile the link between natural resources and sustainable growth may be. Even when large profits were generated and an income re-distribution took place, there appear clear symptoms of natural resource curse in the economy, for them, characterized by an over-appropriation of rents and a systematic deterioration of the quality of policies. Thus, twin surpluses disappeared, monetary policy left the auto-insurance strategy; fiscal policy became strongly pro-cyclical; energy and transportation subsidies grew exponentially leading to an energy deficit and exceeding revenues from export duties; the tax burden rose along with distortionary mechanisms operating with inflation.

Our aim is to empirically study a key relationship to understand the effects of commodity prices on the Argentine economy. The first chain link of effects requires comprehending the effect of prices over both exports and the exchange rate, which are simultaneously related. Therefore, this is the first attempt to econometrically study these transmission mechanisms of world commodity prices. Most of the empirical studies have econometrically modelled exports or exchange rates by using single equations. Those focussed on the effect of commodity prices on economic growth were based on cross-section or panel data (e.g. Collier and Goderis, 2007;

¹Centre of International Economics - Ministry of Foreign Affairs.

Lederman and Xu, 2009; van der Ploeg and Poelhekke, 2009, among others). Several studies also analysed the effect of commodity price shocks on the growth of developing economies (Deaton and Miller, 1995; Brown and Yücel, 1999, Izquierdo *et al.*, 2008; Medina, 2009). In particular, Izquierdo *et al.* (2008) suggest that one-time increases in commodity prices generate level effects on output in Latin American countries. However, they warn that this fact should not be confused with sustained growth.

Our research differs from earlier empirical works as it develops a case study using a simultaneous approach for time series data of exports and real exchange rate. The economic structure of Argentina has suffered significant changes during the last two decades and represents a challenge for the econometric modeling of these relationships.

The paper is organized as follows: the next section outlines the theoretical framework and discuss the transmission mechanisms of commodity prices on exports and real exchange rate. Section 3 describes the data and analyses the degree of persistent behaviour in the variables and their relationships. Section 4 presents the econometric approach for both testing and estimating the long-run structure and short-run dynamics. Section 5 discusses at some length the specification of the deterministic components and presents the empirical results. A reader not interested in the econometric details can skip Section 4 and 5 without losing track. Section 6 sheds light on the results of the econometric approach and takes a closer look on the effects of commodity prices on exports and real exchange rate both in the long-run and short-run. Finally, Section 7 draws our main conclusions.

2 The theoretical background

The framework we adopt is one of a small economy with a commodity-based export structure. We consider a standard commodity-export model in which exports is the difference between domestic supply and demand of exportable goods taken international prices as given (see Corden and Neary, 1982, Corden, 1984, Arize, 1990, Reinhart, 1995; and for the Argentine case see Ahumada, 1996, Catão and Falcetti, 2002). Therefore, the small open economy (SOE) assumption implies that the economy has no influence on international commodity prices and that commodities produced in the country and abroad are homogenous. Thus, we assume the following functional form,

$$x_t = \alpha_0 + \alpha_1 y_t + \alpha_2 p_t^x + \alpha_3 e_t \quad (1)$$

where x stands for the commodity export volume index, y for the sector production capacity, p^x for the world price of exports and e for the real exchange rate. All variables are in logs. The series employed in our empirical investigation are described in the next section.

It is expected that α_1 , α_2 and α_3 will be positive, exports will rise when: (a) there is an increase in the country's capacity to produce commodities, (b) there is an increase in the world price of commodities which makes their production more profitable and discourages the domestic demand of exportable goods, and (c) there is a depreciation of the real exchange rate, having the same effects as in (b). We consider the case of a small commodity-export economy since Argentina can be considered to be a price taker in many of its commodities exports. That is, commodity prices are assumed to be exogenous. This assumption will be tested later.

The SOE assumption also implies that the real exchange rate is the equilibrating force whenever the (domestic) price of exportable goods should change. In Argentina, periods of growing commodity exports would lead to a large inflow of foreign currency, resulting on a further appreciation of the real exchange rate. Therefore, the fact that commodities exports and real exchange rate may be jointly determined in such cases implies that they should be simultaneously modelled.

Many efforts have been done in empirically modelling the exchange rate behaviour to shocks in fundamental factors (see Frankel and Rose, 1995, and Froot and Rogoff, 1995 for a summary)

such as productivity, government spending and the relative price of non-tradables. Empirical studies differ in their choices of underlying real exchange rate fundamentals depending on data availability and/or the analysed economies.

As in Chen and Rogoff (2003), in the Argentina’s case, world commodity price movements (exogenous to this small country) potentially explain exchange rate fluctuations because primary commodities have a significant weight on their trade accounts. Understanding the effects of commodity price shocks on exchange rate is of considerable interest to Argentina, particularly since 2002 after a ten year convertibility regime was abandoned. Being a commodity-export country, an improvement in the terms of trade should tend to appreciate the real exchange rates in line with the hypothesis of the Dutch Disease.

Furthermore, an increase in the domestic consumption (the sum of domestic private and public expenditure) may produce an appreciation of the real exchange rate as the real exchange rate defines the relative price of tradables to non-tradables. Using quarterly data, Rogoff (1992) found that government spending appears to be highly correlated with the real yen/dollar rate, but it does not enter significantly into the regressions once one controls for shocks to the world price of oil. De Gregorio, Giovanni and Wolf (1994) also found that government spending is highly significant for OECD countries. De Gregorio and Wolf (1994) have extended this analysis to incorporate term of trade shocks which were found to be important empirically, though productivity and government spending differentials continue to be important.

In this context, we consider the following model that relates the real exchange rate with real economic fundamentals in a commodity-export country,

$$e_t = \beta_0 + \beta_1 d_t + \beta_2 yprod_t + \beta_3 p_t^x + \beta_4 x_t \quad (2)$$

As in Edwards (1988), we consider only the real factors that influence the (steady state) “equilibrium exchange rate”. Thus, the real exchange rate (e) depends on domestic consumption (d), commodity sector productivity ($yprod$), world commodity prices (p^x) and commodity exported volumes (x). An increase in private or government expenditure raises the relative price of non-traded goods, appreciating the real exchange rate reflecting a higher demand for non-tradable goods over their supply. An increase in world commodity prices improves the current account as commodity exports become more competitive. Therefore, higher commodity terms of trade tend to appreciate the real exchange rate through income or wealth effects. A rise in the commodity sector productivity may raise the relative price of non-tradable goods and appreciate the real exchange rate as the productivity increase is biased towards tradable goods. This may be indicative of the Balassa-Samuelson effect. However, the empirical evidence in favour of such effect is weaker than commonly believed (Froot and Rogoff, 1995).²

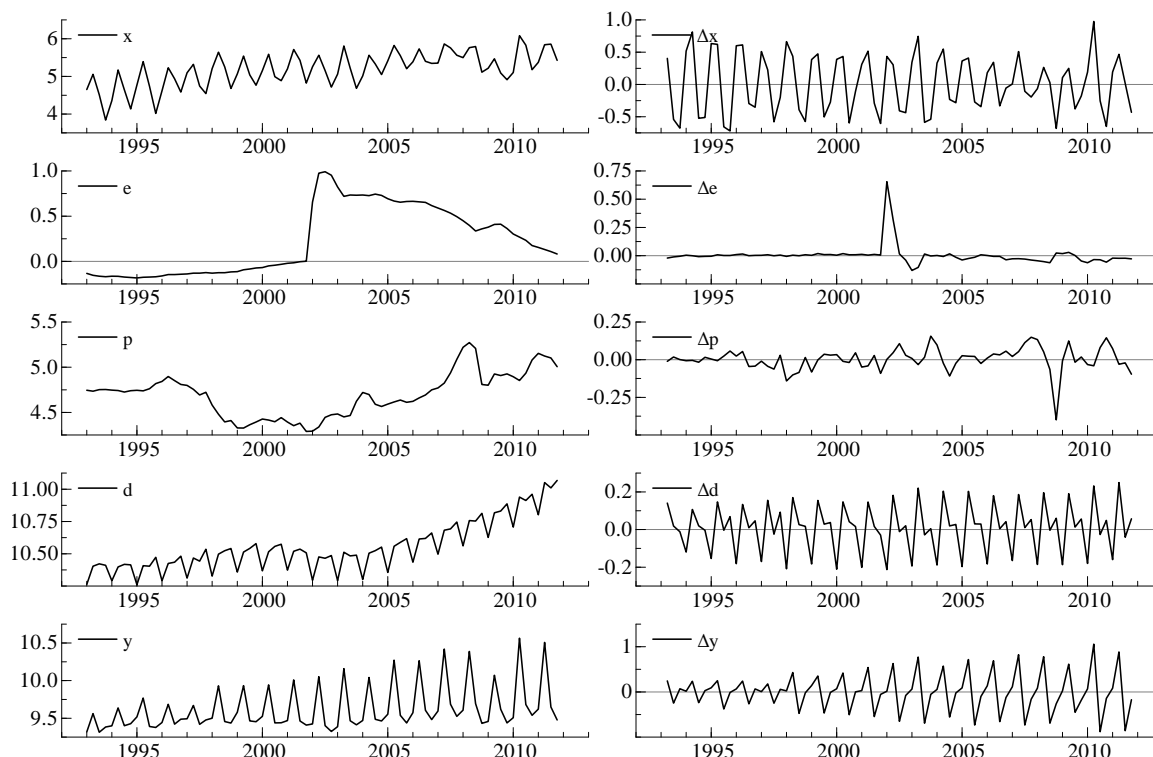
In this paper, we empirically assess the existence of the two long-run (cointegration) relationships expressed in Equation (1) and (2) by applying Johansen (1996) maximum-likelihood cointegration system approach, as described in the Section 4. This approach allows us to jointly estimate these equations as long run relationships of integrated variables. Therefore our econometric study starts with the estimation of a general model which encompasses the variables that enter both the excess export supply function (1) and the real exchange rate determination function (2). After finding evidence about these equations for the Argentine case, an equilibrium correction representation is also estimated to allow for short-run behaviour since we are also interested in studying commodity exports and exchange rate dynamics in a small economy.

²In the empirical section we include in the close system the agricultural output. We assume that the expansion of this sector in Argentina over the last decade may partially depend on its increasing productivity.

3 Data description

The data, which are shown in Figure 1, are quarterly, over 1993Q1 to 2011Q4 ($T=76$). The index of commodity prices³ (p^x) and the real peso/dollar exchange rate⁴ (e) are obtained from the Central Bank of Argentina (both average quarterly). The agricultural sector GDP (y) and the domestic consumption (d), calculated as the sum of public and private expenditures (at constant prices), as well as the raw material export volume index (x) are obtained from the Argentine National Institute of Statistics.

Figure 1: The data (in logs and log-differences)



From Figure 1 we can distinguish two periods according to the evolution of the real dollar/peso exchange rate. The first period (1993-2001) was characterised by the convertibility regime which backed the monetary base with external reserves to guarantee the one peso to one dollar rate of exchange. At the beginning of the sample, the traditional agricultural exports and the agricultural sector grew at moderate rates, particularly soon after the 1995 recession followed by the Tequila crisis. By that time, domestic consumption started to grow at higher rates but it declined deeply in the 2002 recession. The economic recession at the end of the convertibility regime (1998-1999) coincides with declining world commodity prices.

The exchange rate regime finally collapsed in January 2002, after the government announced the default on its sovereign debt and the abandonment of the convertibility. The real exchange rate jumped 93% on a quarterly basis. Argentina's impressive recovery since 2002 coincides with a period of historically high commodity prices. The new scenario for commodity prices has had a great impact on many commodity-export economies, such as Argentina. We can observe that the growth rate of domestic consumption (private and public consumption) has been intensified since 2003 when demand policies also took place. It is worth noting that the world crisis started

³The index of commodity prices (IPMP) developed by the Central Bank of Argentina includes the prices of the most representative commodities for Argentine exports, updating the weights every year to better reflect the product share in Argentine trade.

⁴World commodity prices are measured in US dollars and thus, the value of commodity exports is expressed in US dollars.

in 2008 affected Argentina through commodity prices, but not through financial restrictions as no large volumes of new debt had been acquired after the default. However, expansionary fiscal and monetary policies stimulated economic growth. During 2010 and 2011, output grew at high rates again and exports and imports reached record levels.

Regarding the different components of the series, Figure 1 shows a marked seasonal pattern of commodity exports, agricultural production and domestic consumption. In particular, agricultural production exhibits a changing seasonality which is higher over time suggesting that the agricultural sector becomes more dependent on seasonal effects (see Appendix A for seasonally adjusted figures).

To analyse the degree of persistent behaviour in the variables we conduct univariate unit-root tests as reported in Table 1. We use the ADF (Augmented Dickey-Fuller) test to examine the order of integration of the original variables and their changes. Results indicate that the (log) level of agricultural exports, real exchange rate, commodities prices and agricultural GDP appear to be I(1).

Table 1: ADF statistics for testing unit root (Quarterly Data, 1993-2011)

Sample period 1993Q1-2011Q4									
Variable	k	$t_{ADF(k)}$	ρ	σ	t -prob	AIC	Constant	Trend	Seasonals
x	2	-3.336*	0.540	0.172	0.045	-3.414	Yes	Yes	Yes
x	2	-1.839	0.865	0.181	0.001	-3.329	Yes	No	Yes
e	1	-1.591	0.945	0.083	0.000	-4.912	Yes	Yes	No
e	1	-1.813	0.952	0.083	0.000	-4.939	Yes	No	No
d	4	0.298	1.012	0.022	0.001	-7.506	Yes	Yes	Yes
d	4	1.756	1.043	0.022	0.001	-7.517	Yes	No	Yes
p^x	2	-1.854	0.923	0.073	0.116	-5.170	Yes	Yes	No
p^x	2	-1.285	0.952	0.074	0.119	-5.161	Yes	No	No
y	4	-3.427*	0.196	0.088	0.004	-4.738	Yes	Yes	Yes
y	4	-1.550	0.840	0.093	0.033	-4.626	Yes	No	Yes
Δx	1	-9.535***	-0.573	0.185	0.000	-3.278	Yes	Yes	Yes
Δx	1	-9.618***	-0.570	0.184	0.000	-3.305	Yes	No	Yes
Δe	1	-5.334***	0.307	0.083	0.113	-4.913	Yes	Yes	No
Δe	1	-5.231***	0.331	0.083	0.133	-4.925	Yes	No	Yes
Δd	4	-2.282	0.216	0.022	0.152	-7.539	Yes	Yes	Yes
Δd	4	-1.501	0.598	0.022	0.079	-7.519	Yes	No	No
Δp^x	1	-5.988***	0.150	0.074	0.040	-5.147	Yes	Yes	Yes
Δp^x	1	-5.952***	0.173	0.074	0.048	-5.165	Yes	No	Yes
Δy	3	-5.420***	-1.508	0.095	0.051	-4.588	Yes	Yes	No
Δy	3	-5.462***	-1.508	0.094	0.049	-4.616	Yes	No	No
$\Delta^2 d$	3	-9.248***	-3.803	0.022	0.019	-7.485	Yes	Yes	Yes
$\Delta^2 d$	3	-9.322***	-3.787	0.022	0.019	-7.511	Yes	No	Yes

Notes: k indicates selected lag-length that minimizes the Akaike Information Criterion (AIC). The columns report the name of the variable examined, the selected lag length (k), the ADF statistic ($t_{ADF(k)}$), the estimated coefficient on the lagged level that is being tested for a unit value (ρ), the regression's residual standard error (σ), the tail probability of the t-statistic on the longest lag of the final regression (t -prob), the AIC and the columns indicating the included deterministic components. *, ** and *** indicate significance at the 10%, 5% and 1% level, respectively.

Table 1 also suggests that the domestic consumption is I(2) as the log-difference has a unit root and its second-difference is stationary. The fact that the domestic consumption could be considered as an integrated variable of order 2 implies that there is still great persistence

remaining in the growth rate of this variable (as can be appreciated in the seasonally adjusted data of Appendix A). That is, domestic consumption seems to have a changing growth rate since the mid-2000s, when fiscal policy became strongly pro-cyclical. This may be indicative of an I(2) behaviour, particularly for the last years of the sample. Although this variable would be the only I(2), its effect on a system estimation should be considered. In particular, a typical smooth behaviour of a stochastic I(2) trend can often be approximated with an I(1) stochastic trend around a broken linear deterministic trend (see Juselius, 2006: p. 294).

An appropriate specification of the deterministic components, such as trends, constant and dummies, and how they will enter the model is an important issue to be considered in the empirical modelling. This is because the chosen specification is likely to strongly affect the reliability of the model estimates and to change the asymptotic distribution of the cointegration test. Argentina's economy has suffered important economic crisis and changes in its economic structure in the last two decades. Different exchange rate regimes as well as fiscal and monetary policy changes represent a challenge to econometrically estimate economic relations in this country. Therefore, we carefully study the different breaks and shifts in the considered series. We discussed these components in the following sections and in Appendix A.

Furthermore, we test for multivariate stationarity, that is, trend stationarity of the variables entering the system. These results are shown in Section 5, after the cointegration analysis. According to these statistics and by including a linear broken trend, a I(1) system was estimated to study cointegration, as shown in the next section. A reader not interested in the empirical/econometric methodology can jump to Section 6 without losing track.

4 The econometric approach

We consider a 5-dimensional VAR model for $z_t' = [x_t, e_t, d_t, p_t^x, y_t]$, where x_t stands for natural resources export volume, e_t for real exchange rate, d_t for real domestic consumption, p_t^x for real commodity prices, and y_t for agricultural sector gross domestic product. Small letters denote their logarithmic values. From the estimation of this system we try to identify cointegration relationships which may represent the long run equilibria from the economic structure. A great advantage of this approach is the invariance of the cointegration property to extension of the information set (see Juselius, 2006, chapter 19). This means that once cointegration is found among a set of variables, the cointegration results will remain valid if more variables are added to the model. Therefore, no "omitted variable effects" are present for cointegration when we adopt a specific-to-general strategy.⁵ The model is structured around r cointegration relations (the endogenous or pulling forces) corresponding to $p-r$ stochastic trends (the exogenous or pushing forces). Therefore, the pulling force is formulated as a dynamic adjustment model in growth rates and equilibrium errors, the so-called Vector Equilibrium Correction Model (VEqCM),

$$\Delta z_t = \alpha\beta' z_{t-1} + \Gamma_1 \Delta z_{t-k} + \Phi D_t + \varepsilon_t \quad (3)$$

where z_t is a p -dimensional vector of economic variables, D_t is a $m \times 1$ vector of m deterministic terms, $\varepsilon_t \sim Niid(0, \Omega)$ is a $p \times 1$ vector of errors, Δ is the first difference operator, α , β are $p \times r$ coefficient matrices, Γ_1 is a $p \times p$ matrix of short-run adjustments coefficients, Φ is a $p \times m$ matrix of coefficients, and the lag length k in the corresponding VAR.

Such model is designed to distinguish between influences that move equilibria (pushing forces) and influences that correct deviations from equilibrium (pulling forces) which give rise to long-run relations (see Juselius, 2006). The division between pulling and pushing forces is based on the cointegration rank, r , imposed as a reduced rank restriction in the VAR model.⁶

⁵However, the common trend representation is not invariant to changes in the information set as they are the residuals of the VAR used for the impulse-response analysis.

⁶Tests of the hypothesis of r cointegrations vectors can be based on the trace statistics.

Therefore, after determining the cointegration rank, the r -column vectors of β (the eigenvectors) allow us to find the long run solutions of economic models. But cointegration by itself does not imply which variable adjust to reach the equilibrium. The coefficients in α give the information about which variables adjust and thereby weak exogeneity can be tested by zero restrictions in the respective coefficient, as suggested by Johansen (1992) and Urbain (1992).

Finally, it is important to note that the inclusion of deterministic components (trend, different kind of dummies for the whole period and sub-periods) in the models is critical for the rank determination. In the case of the Argentine series, the choice of deterministic components is particularly a difficult task. The following sections focus particularly on analysing the appropriate deterministic components to be included in the system.

5 Empirical Results

This section first presents cointegration results when considering a broken linear trend and other deterministic components (Section 5.1), it then shows the dynamic estimations from a simultaneous equilibrium correction model (Section 5.2). The purpose of this section is, therefore, to discuss how we applied the econometric approach.

5.1 Cointegration analysis

To estimate the 5-dimensional VAR model from the data previously described, based on prior knowledge of relevant historical events and time properties of the series, we introduced the following deterministic components: centred seasonal dummies to control for the observed pronounced seasonal pattern in x , d and y ; two dummy variables for 2002Q1 (Argentine economic crisis) and 2008Q4 (worldwide fall in commodity prices); a linear trend and a broken linear trend for the period 2002Q1-2011Q4. Dummies are included unrestrictedly in the system. Both trends are restricted to enter the cointegration space since the variables can cointegrate but have different deterministic trends (see Juselius, 2006, p.98). In the case we analysed they may show also deterministic trend differences after 2002 when a new economic regime started and many variables steady grew (see Appendix A). A step dummy 2002Q1-2011Q4 was also incorporated (unrestrictedly) to allow the growth rates to change due to similarity conditions of the broken trend, as suggested by Nielsen and Rahbek (2000).

After having account for extraordinary events over the sample period, the information criteria suggest different values of k , in which case we prefer the Schwarz criterion for selecting the most parsimonious model with $k=2$ lags (Juselius, 2006: p.72).⁷ Despite the instability of the sample after including the deterministic components in the VAR, as before described, normality tests, reported in Table 2, show that multivariate normality is not rejected.⁸ However, there is a lack of normality in some specific variables: the real exchange rate at a 10% level and the agricultural production mainly due to excess kurtosis.

⁷Adding too many lags is more harmful for the results than accepting some moderate residual autocorrelation in the model. This is because regime shifts, non-constant parameters, etc. are often difficult to diagnose in a heavily over-parametrized model. Furthermore, residual autocorrelation in a first tentative VAR(2) model is more often associated with structural misspecification, rather than with left-out dynamics.

⁸We use the Doornik and Hansen (1994) multivariate test to test residual normality in a VAR system.

Table 2: Specification tests for the unrestricted VAR(2) model (p -values)

Equation	Skewness	Kurtosis	Sk+Kr
x	0.07	2.55	1.52 (0.47)
e	-0.08	3.95	5.64 (0.06)
d	-0.03	2.94	0.60 (0.74)
p^x	-0.02	2.98	0.61 (0.74)
y	-0.09	3.98	11.93 (0.00)
System			12.26 (0.27)

Because the asymptotic distribution for the rank test depends on the deterministic components included in the model, we followed Johansen, Mosconi and Nielsen (2000) to empirically test cointegration in the presence of a broken linear trend. We compute critical values using the response surface function from the Monte Carlo study of theirs. The correct choice of the cointegration rank, r , will influence all subsequent econometric analysis and may very well be crucial for whether or not we reject our prior economic hypotheses (Juselius, 2006: p.140).

Table 3 reports the estimates eigenvalues, λ_i and the 95th percentile of the Γ -distribution when considering a broken trend and a shift dummy in the cointegration relations, $C_{.95}$ (see Nielsen, 1997 and Doornik, 1998).

Table 3: The rank test of cointegration

r	$p-r$	l	λ_i	$Test$	$C_{.95}$
0	5	1	0.84	244.13	119.21
1	4	2	0.60	107.93	89.03
2	3	3	0.29	39.70	62.50
3	2	4	0.11	14.80	39.77
4	1	5	0.08	5.84	20.26

Therefore, the tests for cointegration rank supports $r=2$. That is, two cointegration vectors can be obtained.

Table 4 reports values of the multivariate statistic for testing trend stationarity of the variables entering the system. Specifically, this statistic tests the restriction that the cointegrating vector contains all zeros except for a unity corresponding to the designated variable and an unrestricted coefficient of the trend and broken trend.

Table 4: Multivariate stationarity test

	x	e	d	p^x	y
$\chi^2(3)$	49.95***	39.04***	60.84***	53.65***	41.59***

All tests reject the null of stationarity. By being multivariate, these statistics may have higher power than their univariate counterparts. Also, the null hypothesis is the stationarity of a given variable rather than the nonstationarity thereof, and stationarity may be a more appealing null hypothesis. That said, these rejections of stationarity are in line with the inability to reject in Table 1 the null hypothesis of unit root in each of e , x , p^x , d and y .

Having determined the cointegration rank, the cointegrated VAR can be estimated and the dimensions of the α and β matrices can also be determined. Table 5 shows the adjustment coefficients, α , and the identified eigenvectors, β (once restrictions on α have been imposed). The first two columns show the unrestricted adjustment coefficients. Although the adjustment of the exchange rate appears as significant in vector 1, once restrictions on the not significant adjustments coefficients (but not on the eigenvectors) are imposed, the α of the exchange rate is no longer significant. Therefore, we also tested that restriction. Finding weakly exogenous variables by testing the hypothesis that certain variables do not adjust to long-run relations is helpful to identify the common driving trends and the long-run structure. Therefore the

second two columns report the restricted adjustment coefficients and identified eigenvectors when normalizing by one element in each vector (x and e , respectively) since it is only possible to get standard errors of β when each cointegration vector is properly normalized. The last two columns report the resulting restricted cointegration relations.

Table 5: Cointegration vectors

Adjustment coefficients α																
r	(1)		(2)		(1)		(2)		(1)		(2)					
Variable	Coef	SE	Coef	SE	Coef	SE	Coef	SE	Coef	SE	Coef	SE				
x	-0.54	0.08	-0.31	0.38	-0.55	0.08	0	0	-0.57	0.08	0	0				
e	0.03	0.01	-0.34	0.04	0	0	-0.38	0.04	0	0	-0.35	0.04				
d	0.01	0.01	-0.14	0.06	0	0	-0.14	0.06	0	0	-0.14	0.05				
p^x	0.01	0.03	-0.07	0.14	0	0	0	0	0	0	0	0				
y	0.42	0.06	0.14	0.27	0.43	0.05	0	0	0.44	0.05	0	0				
$\chi^2(j)$						0.46										
p -value						0.98										
j						4										
Eigenvectors β																
r	(1)				(2)				(1)				(2)			
Variable	Coef		SE		Coef		SE		Coef		SE		Coef		SE	
x	1		0		0.01		0.03		1		0		0		0	
e	-0.68		0.29		1		0		-0.54		0.28		1		0	
d	-0.78		0.54		1.39		0.20		0		0		1.54		0.12	
p^x	-0.39		0.13		0.20		0.05		-0.33		0.12		0.23		0.04	
y	-2.63		0.16		0.07		0.07		-2.58		0.16		0		0	
tr	-0.01		0.00		-0.01		0.02		-0.01		0.00		-0.01		0.00	
tr_{02-11}	0.02		0.01		0.00		0.00		0.02		0.01		0		0	
$\chi^2(j)$						4.83										
p -value						0.78										
j						8										

Notes: other unrestricted variables included are cseasonals, a step dummy for 2002Q1-2011Q4 and dummies for: 2002Q1 (peso devaluation) and 2008Q (commodity crisis).

An important part of a long-run cointegration analysis is to test (over-) identifying restrictions on β to achieve economic interpretability. Table 5 also reports long-run weak exogeneity results for both the restricted and unrestricted β coefficients which allow us to estimate an export model and an exchange rate model at traditional significance levels. As the agricultural GDP and the domestic demand variables also adjust in vector 1 and 2, respectively, the equations can be re-parameterized by normalizing on these variables and still have an economic interpretation. For the system as a whole, commodity prices were empirically detected as weakly exogenous, that is, commodity prices have influenced the long-run stochastic path of the other variables in the system, while at the same time has not been influenced by them.

This result validates the small country assumption and the price-taking hypothesis holds as the commodity price index (being an export-weighted average of individual prices) is found to be weakly exogenous. We tested this assumption because the increasing Argentine participation on soybean trade may suggest that commodity prices are not exogenous for the commodity-export model of Argentina. Furthermore, in this study we use an export-weighted commodity price index which might have implied that commodity prices were not exogenously determined in the export-determination model.

In the next graphs we can observe the stability of the cointegration results, reported in Table 5, according to the recursive estimation of the system. First, we can see the recursive eigenvalues in Figure 2(a). Although there might be a third vector in the first part of the sample, it was not significant for the whole period. Given the imposed restrictions, the estimates of the effects of

the explanatory variables in both vectors remained constant for the last part of the sample, as shown in Figure 2(b). We found some changes in coefficient estimates during 2002/2003 when the economy began to recover after a deep recession from 1999 to 2001.

Figure 2a: Recursive eigenvalues

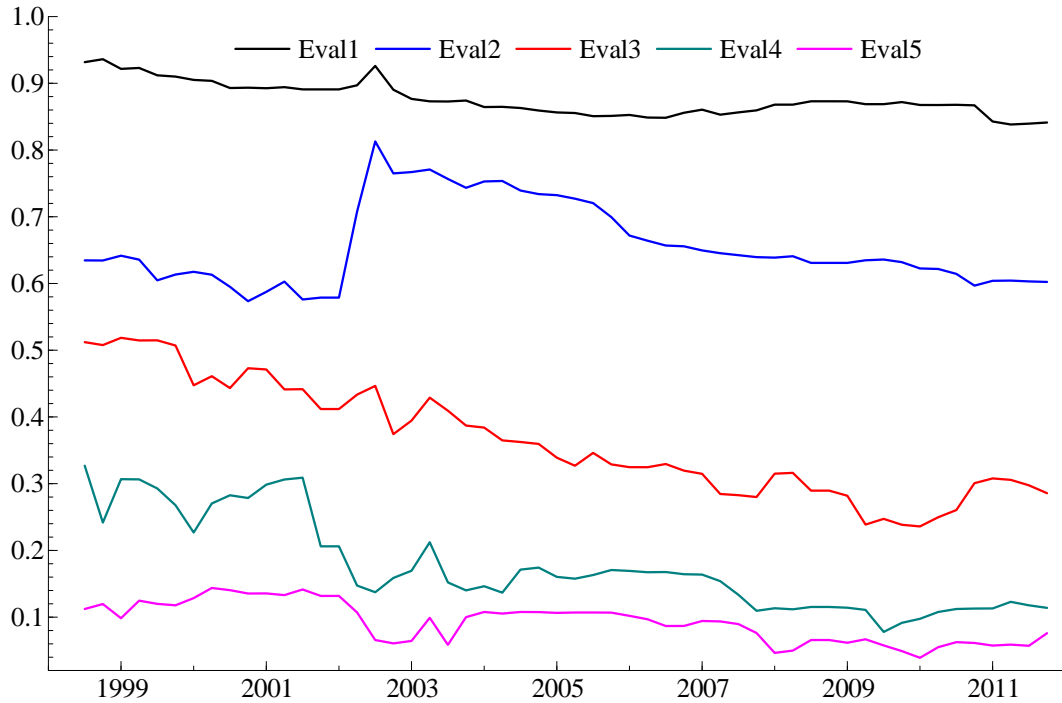
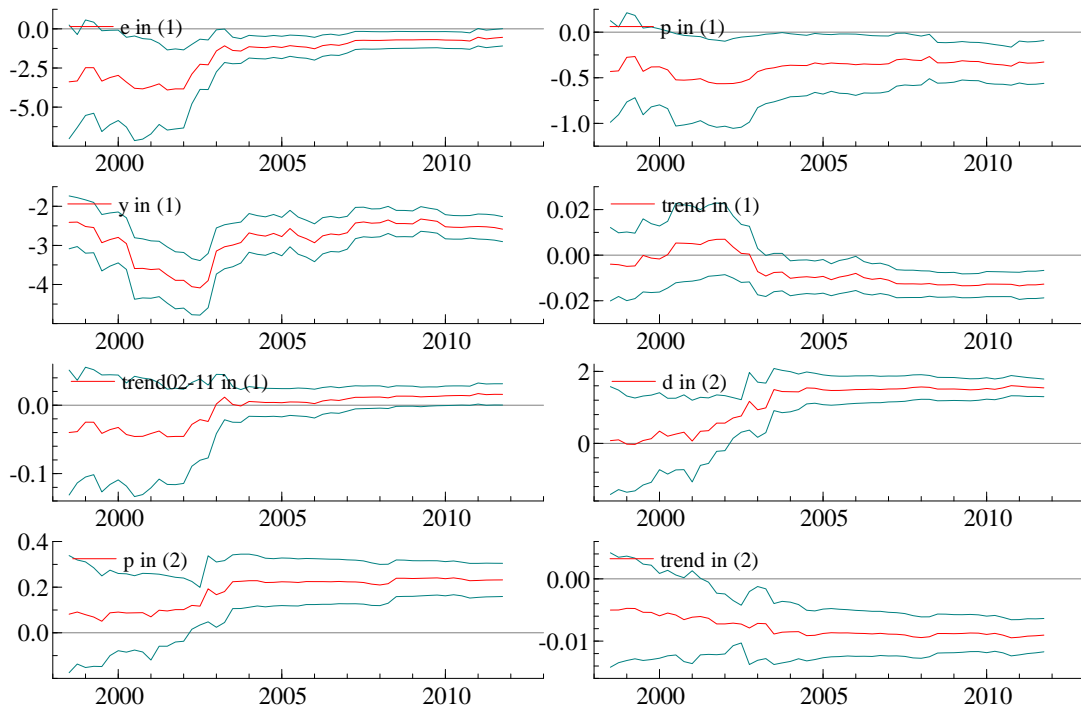


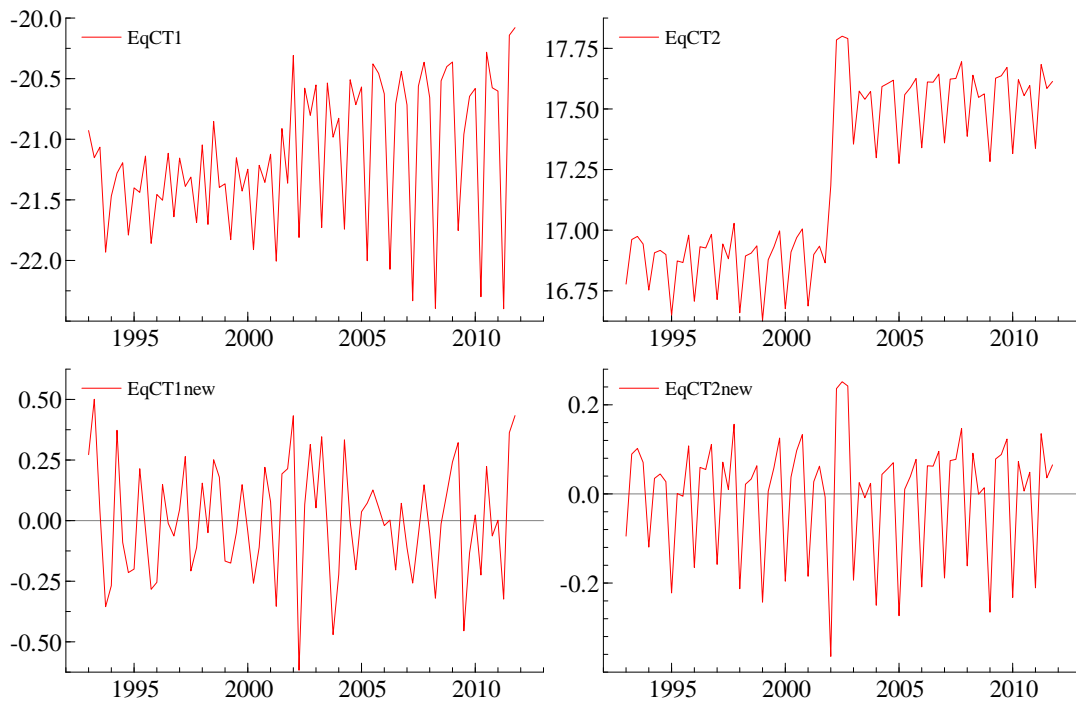
Figure 2b: Recursive eigenvectors ($\beta \times \pm 2SE$)



5.2 A simultaneous equilibrium-correction model

Following the proposed normalization of the two identified vectors (as shown in Table 5), we then estimate a vector equilibrium correction model (VEqCM) that considers the simultaneously correction of the variables. The analysis of the last section follows the standard literature about the use of deterministic components, but as the next figures suggest we need adjust the mean of the equilibrium correction terms additionally. Figure 3 illustrates the behaviour of both EqC terms (EqCT) which show that the first cointegration relation (EqCT1) has a stochastic seasonal pattern and the second vector (EqCT2) has a mean change since 2002 reflecting the abandonment of the convertibility regime. The seasonal pattern of EqCT1 is stochastic, that is, the seasonal effect becomes higher over time suggesting that commodity exports and agricultural sector becomes more dependent on seasonal factors. Therefore, EqCT1 was corrected by obtaining the residuals after regressing it on centred seasonals, broken centred seasonals for 2003-2011 and a step dummy for the period 2002-2011, whereas EqCT2 is regressed on a step dummy for the period 2002-2011. The corrected EqCT1 and EqCT2 are denoted by EqCT1new and EqCT2new and they show a stationary pattern at the bottom of Figure 3.

Figure 3: Equilibrium correction terms from vector 1 and 2



Using the identified cointegration relations, we first estimated a multivariate dynamic equilibrium correction model for the system. By first removing insignificant lagged variables of the system and then removing insignificant coefficients from the equations, based on a Likelihood Ratio test, we arrived at the parsimonious model presented in Table 6. Therefore, a restricted simultaneous EqCM is estimated by Full Information Maximum Likelihood (FIML). The column heading in the top of the table indicates the dependent variable in each of the model equations, while the row headings indicate the conditioning variables.

Table 6: FIML Estimation of the EqCM

Variable	Δx_t	Δy_t	Δe_t	Δd_t
<i>constant</i>	0.01 (0.02)	0.03 (0.01)	0.001 (0.002)	0.002 (0.003)
<i>EqCT1new_{t-1}</i>	-0.23 (0.09)	0.10 (0.04)		
<i>EqCT2new_{t-1}</i>			-0.16 (0.05)	-0.40 (0.08)
Δx_{t-2}	-0.42 (0.10)			
Δy_{t-1}	0.24 (0.13)	-0.53 (0.08)		
Δy_{t-2}	0.31 (0.12)	-0.65 (0.07)		
Δy_{t-3}		-0.49 (0.09)		
Δe_t				-0.42 (0.16)
Δe_{t-1}			0.42 (0.03)	0.17 (0.07)
Δe_{t-2}			-0.15 (0.03)	
Δe_{t-3}				0.07 (0.03)
Δe_{t-4}	0.40 (0.18)		-0.13 (0.02)	
Δd_{t-1}		-1.15 (0.30)		
Δd_{t-2}			-0.30 (0.07)	
Δd_{t-3}			-0.37 (0.08)	
Δd_{t-4}			-0.27 (0.08)	0.22 (0.12)
Δp_t^x			-0.07 (0.02)	
Δp_{t-2}^x				0.13 (0.04)
$\widehat{\Sigma}$ (standard errors on the diagonal, off-diagonal elements are correlations)				
Δx_t	0.14			
Δy_t	0.46	0.06		
Δe_t	-0.18	-0.30	0.01	
Δd_t	0.02	-0.03	0.19	0.02
AR(1-5)	1.22 [0.31]	0.70 [0.63]	1.74 [0.14]	2.51 [0.04]
ARCH(1-4)	0.39 [0.82]	1.07 [0.38]	0.83 [0.51]	0.70 [0.60]
Normality: $\chi^2(2)$	1.07 [0.59]	5.76 [0.06]	0.71 [0.70]	0.48 [0.79]
Heteroskedasticity	0.69 [0.78]	1.34 [0.21]	0.81 [0.70]	0.72 [0.81]
Vector SEM-AR: F(80,136): 1.06 [0.38]				
Vector Normality: $\chi^2(8)$: 11.53 [0.17]				
Vector ZHetero: F(204,54): 1.07 [0.39]				

Notes: standard errors are reported in parentheses and p-values in brackets. The following deterministic components were included: centred seasonal dummies (cseasonals), broken cseasonal for the second quarter during 2003-2011, and dummies for 2002Q1 (peso devaluation), 2009Q2 (the worst drought crisis in 100 years) and 2010Q2 (recovery after drought).

From Table 6 we can observe that the adjustment coefficients (the equilibrium-correction term, EqCT) are significant and have the expected sign.

Since we adopt a simultaneous modelling approach we can test contemporaneous effects. In particular, the agricultural production growth (Δy) and commodity-export growth (Δx) have

the highest residual correlation as observed in Table 6, while the others are much lower. No significant effects were detected for these variables. In terms of the contemporaneous effects, the growth of commodity prices (Δp) and the depreciation rate (Δe) are significant.

6 Discussing long-run and short-run effects on exports and real exchange rate

This section discusses the empirical results presented in last section. We divide the economic interpretation into two main sub-sections discussing the long-run effects (6.1) and the short-run effect (6.2) on exports and real exchange rate.

6.1 Long-run relations: pulling and pushing forces

Given the above results and the tests conducted in Section 5, we identify two long-run (cointegration) relations. The long run equations are,

$$x_t = constant + \underset{(0.27)}{0.54}e_t + \underset{(0.12)}{0.33}p_t^x + \underset{(0.16)}{2.58}y_t + \underset{(0.003)}{0.01}trend - \underset{(0.01)}{0.02}trend_{02-11} \quad (4)$$

$$e_t = constant - \underset{(0.12)}{1.54}d_t - \underset{(0.04)}{0.23}p_t^x + \underset{(0.001)}{0.01}trend \quad (5)$$

The estimated cointegration relations show the factors affecting both natural resources exports and real exchange rate in the long run. The first relation (equation 4) describes the factors affecting commodity exports, while the second relation (equation 5) describes a model of real factors affecting the real exchange rate.

From Equation (4), we can observe that natural resources exports positively depend on the exchange rate, that is, a depreciation of real exchange rate leads to an increase in the exported quantity of raw materials as Argentine commodities exports become more competitive. Also a significant and positive effect of real commodity prices on exports is found. An increase in commodity prices will encourage commodity exports.

Exports are more elastic to variation in the real exchange rate (0.54) than in the world price (0.33). Furthermore, we found that an increase of the agricultural sector GDP raises commodity export supply (the elasticity is 2.54). We also detected that a linear trend is significant in both equations. This indicates different linear trends of the variables entering our model, that is, the variables grow at different rates over the sample. For the equation, such differences also changed since 2002 as the broken trend ($trend_{02-11}$) is significant too.

From Equation (5), the real exchange rate negatively depends on domestic consumption. An increase in domestic consumption (say 1%) implies an increase of non-tradable good prices pushing down the real exchange rate (in near 1.5%). These results reflect a higher demand for non-tradable goods over their supply during the sample. We also found that a 10% increase in international commodities prices leads to an approximate 2.3% appreciation of the real exchange rate. This finding is in line with the hypothesis of Dutch Disease in which an improvement in the terms of trade tends to appreciate the real exchange rate. No evidence of Balassa-Samuelson effect was found, when we used the agricultural production as an (imperfect) proxy of trade sector productivity.

The econometric analysis in Section 5 allowed us to test exogeneity, rather than assuming from the outset which variables are exogeneous and which not. We found that for the system as a whole, only commodity prices are weakly exogenous (the pushing variables), that is, world commodity prices have influenced the long-run stochastic path of the other variables, and not viceversa. This result validates the small country assumption and the price-taking hypothesis holds. But which are the equilibrium correcting variables in each long-run relation? We find out

that in the first relation both the commodity exports and the agricultural production adjust to a disequilibrium from the steady-state. For the second relation, both the real exchange rate and the domestic consumption adjust to correct the deviations from the long run. Based on these results, we jointly estimated the equations presented in 6.2.

6.2 Dynamics in commodity exports and real exchange rate

By allowing for dynamic adjustment towards long-run steady states, we were able to estimate the short-run effects for the four variables which are simultaneously determined: commodity exports (x), agricultural production (y), real exchange rate (e) and the domestic consumption (d). These short-run equations are jointly estimated by Full Information Maximum Likelihood (FIML, where Δ denotes quarterly growth and the EqCT indicates the deviations from the long-run equilibrium).

We first analyse the estimated commodity exports growth equation,

$$\Delta x_t = 0.01 - 0.23 EqCT1new_{t-1} - 0.42 \Delta x_{t-2} + 0.24 \Delta y_{t-1} + 0.31 \Delta y_{t-2} + 0.40 \Delta e_{t-4} \quad (6)$$

(0.02) (0.09) (0.10) (0.13) (0.12) (0.18)

The growth of commodity exports shows an autoregressive structure. The speed of adjustment coefficient suggests that 23% of the disequilibrium (excess export supply) disappears in the first quarter.⁹

Commodity exports growth are explained by the growth of agricultural production, which is significant and it has the expected positive sign indicating that the agricultural domestic supply capacity during the last two-quarters positively impacts on the commodity exports. Furthermore, last year's depreciation rate has also a positive and significant effect on exports growth. Higher depreciation rates and thus, a more competitive real exchange rate, succeed in boosting commodity exports.

As regards the agricultural production dynamics, we estimated the following equation,

$$\Delta y_t = 0.03 + 0.10 EqCT1new_{t-1} - 0.53 \Delta y_{t-1} - 0.65 \Delta y_{t-2} - 0.49 \Delta y_{t-3} - 1.15 \Delta d_{t-1} \quad (7)$$

(0.01) (0.04) (0.08) (0.07) (0.09) (0.30)

The agricultural production growth has also an autoregressive behaviour. The lower coefficient of the equilibrium correct term with respect to that of commodity exports suggests a slower adjustment, that is, the agricultural production growth corrects 10% of the long-run deviations in the first quarter. The positive sign is due to the fact the EqCT is formulated for commodity exports in which agricultural production has a positive effect on exports.

The domestic absorption has a short-run negative effect on agricultural production. This finding suggests that, over the sample, an increase in government and private expenditure may be associated with raising (expectations of) non-tradable prices and the costs of the agricultural sector. Therefore, higher production costs would lower agricultural output.

With respect the dynamics of the real exchange rate we found,

$$\Delta e_t = 0.001 - 0.16 EqCT2new_{t-1} + 0.42 \Delta e_{t-1} - 0.15 \Delta e_{t-2} - 0.13 \Delta e_{t-4} \quad (8)$$

(0.002) (0.05) (0.03) (0.03) (0.02)

$$- 0.30 \Delta d_{t-2} - 0.37 \Delta d_{t-3} - 0.27 \Delta d_{t-4} - 0.07 \Delta p_t^x$$

(0.07) (0.08) (0.08) (0.02)

The depreciation rate is negatively affected by changes in the domestic absorption and in world commodity prices. That is, an increase in domestic consumption (via an expansionary

⁹After the introduction of dummies for changing seasonality, up to four lagged effects are detected and the adjustment coefficients are lower than before in the export and agricultural production equations.

fiscal policy or an increasing private consumption) appreciates the real exchange rate. Furthermore, there is a simultaneous negative effect of commodity prices on real exchange rate. An increase in the world commodity prices that have a significant weight on Argentine trade accounts appreciates the real exchange rate (as in the case of Dutch disease). Empirical evidence of real exchange appreciation during periods of commodity export bonanza was also found by Cashin et al. (2004) for 58 commodity-export countries.

Finally, the domestic consumption dynamics are estimated through the following equation,

$$\begin{aligned} \Delta d_t = & \underset{(0.003)}{0.002} - \underset{(0.08)}{0.40} EqCT2new_{t-1} - \underset{(0.16)}{0.42} \Delta e_t + \underset{(0.07)}{0.17} \Delta e_{t-1} + \underset{(0.03)}{0.07} \Delta e_{t-3} \\ & + \underset{(0.12)}{0.22} \Delta d_{t-4} + \underset{(0.04)}{0.13} \Delta p_{t-2}^x \end{aligned} \quad (9)$$

The domestic consumption adjusts 40% to changes in equilibrium conditions in the first quarter. Domestic absorption growth negatively depends on the exchange rate depreciation and positively on the last two quarters commodity prices growth. Therefore, there is a direct impact of changes in the real exchange rate on domestic absorption, indicating that an increase (a reduction) of the exchange rate tends to depress (stimulate) domestic demand. Changes in world commodity prices of the last two quarters have also a significant, but positive effect of domestic absorption. The commodity prices boom encourages expansionary fiscal policies or private consumption. One of the main sources of the increasing growth of government expenditures has been the revenues from commodity-export duties. Commodity prices have been also affecting consumer decisions through expectations of future income or wealth perception as Ahumada and Garegnani (2012) showed. In their empirical study the soybean price was a useful variable that improves the forecast of a consumption model over the 2006-2011 forecasting period.

From the joint estimation of the last equations we can observe the short run dynamics of commodity exports, agricultural production, real exchange rate and domestic absorption reacting to the direct and indirect effects of commodity prices.

7 Conclusions

In this paper the behaviour of commodity exports and the real exchange rate has been econometrically studied adopting a simultaneous approach to understand how these variables are affected by commodity prices. This is a central issue to analyse for many developing countries whose economies have been deeply transformed by the recent commodity boom. For those facing the positive (and unusual long-lasting) shock on commodity prices, the resulting exports growth of natural resources seems to be a blessing for their economies, but it has had also important consequences derived from the exchange rate appreciation. We develop a case-study for Argentina, a historical well-known commodity producer and exporter. Despite the fact that Argentina has a traditional reputation of being a resource-rich country largely based on agricultural production, the effect of commodity prices on the economy has not been much analysed in empirical research.

From applying Johansen (1996) maximum-likelihood cointegration system approach, we jointly estimate two equations -as long run relationships of integrated variables- which can be identified with the excess export supply function and the real exchange rate determination function under the assumption of a small country. We test exogeneity, rather than assuming from the outset which variables are exogenous and which not. We found that for the system as a whole, only commodity prices are weakly exogenous (the pushing variables) which validates the small country assumption and that the price-taking hypothesis holds.

From the estimated long-run excess export supply function we found that commodity exports positively depend on the exchange rate. A depreciation of the real exchange rate leads to an

increase in the exported quantity of raw materials as Argentine commodities exports become more competitive. Also a significant and positive effect of real commodity prices on exports is found, that is, an increase in commodity prices will encourage commodity exports. Furthermore, from the real exchange determination function we found that the real exchange rate negatively depends on domestic absorption. An increase in domestic consumption implies an increase of non-tradable good prices pushing down the real exchange rate reflecting a higher demand for non-tradable goods over their supply during the sample. We also found that an increase in international commodities prices leads to an appreciation of the real exchange rate. This finding is in line with the hypothesis of Dutch Disease in which an improvement in the terms of trade tends to appreciate the real exchange rate.

As we are also interested in studying commodity exports and exchange rate dynamics, an equilibrium correction representation is also estimated to allow for short-run behaviour.

In a nutshell, the estimated model has shown that commodity prices have been a key variable to explain exports and exchange rate, which are also simultaneously related. Thus, we econometrically modelled the first chain link of the effects of commodity prices on the Argentine economy, which should be extended to empirically study economic growth and income distribution effects. They would be the subject of a future research.

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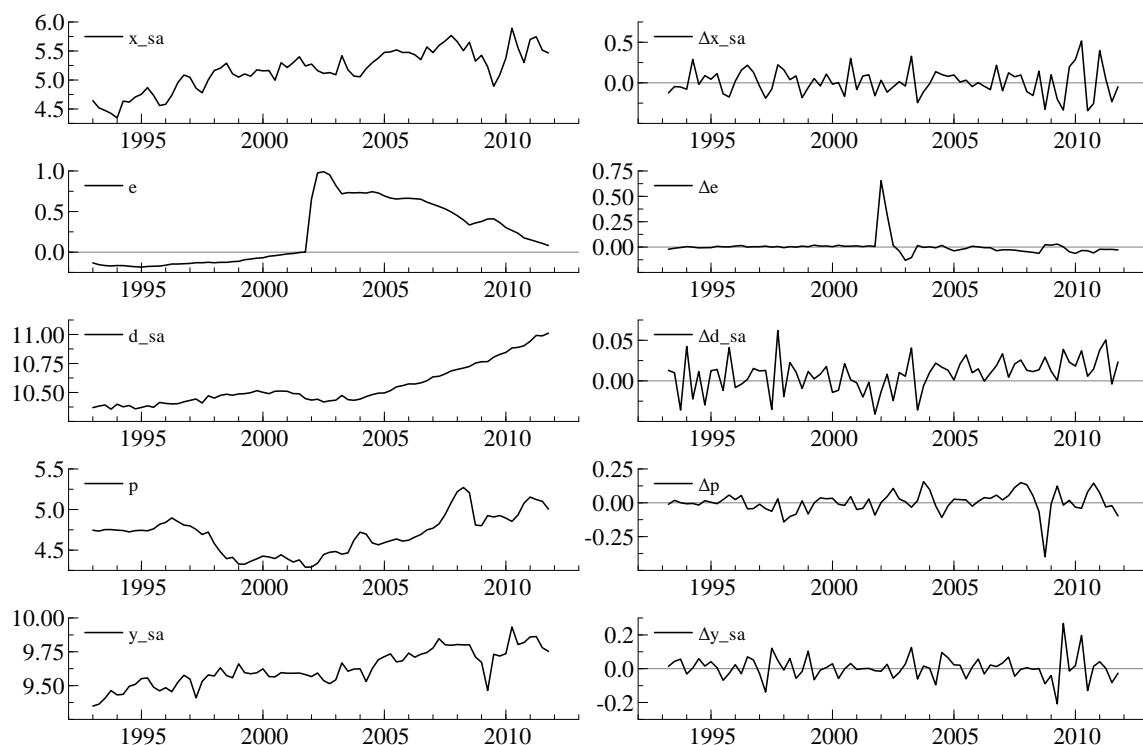
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A Analysis of deterministic components

As explained in Section 4, the domestic consumption appears to be $I(2)$ as the ADF indicates. The graphs of the differenced data are often informative about potential $I(2)$ behaviour, but as our data has a strong seasonality (see Figure 1), we show in Figure A1 the seasonally adjusted variables (*sa* extension) and their first-differences (Δ).¹⁰ From Figure A1 we can observe that the domestic consumption has had a changing growth rate since mid 2000s. During the last decade, fiscal policy became strongly pro-cyclical, the central government has a strongly expansionary fiscal policy.

Figure A1: Seasonally adjusted series (log-level and log-difference)

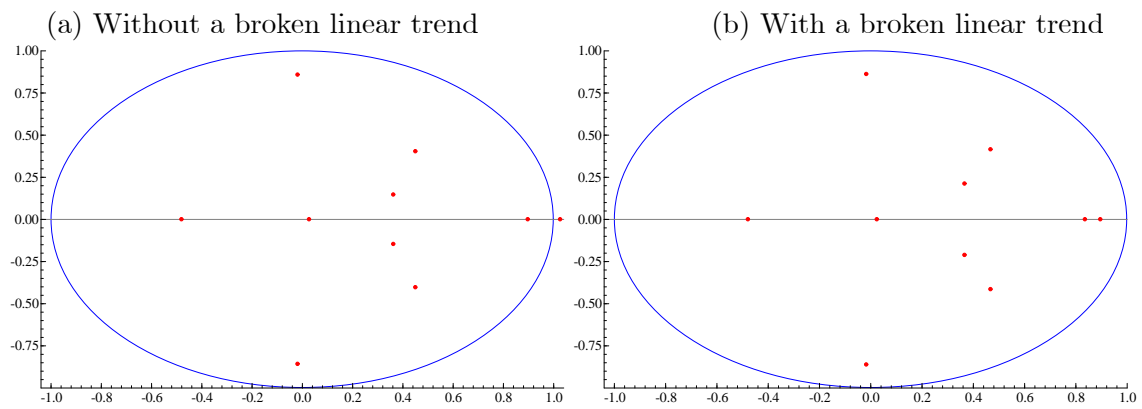


Regarding the system approach, an $I(2)$ analysis can be performed not only when all variables of the system are $I(2)$, but also as an approximation when the acceleration rates become relevant (in addition to growth rates). However, the $I(2)$ analysis may be avoided by allowing for many breaks in the linear trend in an $I(1)$ context. In our case, we included one broken linear trend since 2002 when the different variables started to steady grow, but at different rates, as suggested by Figure A1.

Figure A2 shows the plot of the roots of the companion matrix when the unrestricted VAR(2) is estimated with and without a broken linear trend from 2002Q1 to 2011Q4.

¹⁰Seasonal adjustment was implemented through X12-ARIMA

Figure A2: Root of the companion matrix



From Figure A2 we can notice one root outside the unit circle when the VAR is estimated without a broken linear trend which is a sign of $I(2)$ behaviour in at least one of the variables. However, allowing for a broken linear trend from 2002Q1 to 2011Q4 all the roots of the companion matrix remain inside -but two still are close- the unit circle.