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

Hildegart Ahumada[†] and Magdalena Cornejo[‡]

[†]Di Tella University (hahumada@utdt.edu)
[‡]Buenos Aires University and Di Tella University (mcornejo@utdt.edu)

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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 derived from the real exchange rate appreciation. To take into account the direct and indirect commodity price effects on exports and the 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.

1 Introduction

The commodity boom over the last decade seems to have opened up a new reality for developing economies that produce and export natural resources. For those facing the positive (and unusually long-lasting) shock of high commodity prices, the resulting export growth of natural resources seems to be a blessing for their economies, but it may have also had important side effects on the exchange rate, the economic structure and income distribution. Furthermore, recurrent patterns of commodity booms and busts have created significant uncertainty for commodity exporting economies.

There exists much 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. This has two major effects: a real exchange rate appreciation and an increase of their returns on production relative to other tradable goods. Therefore, there are no incentives to invest in other tradable goods, resulting in a highly commodity-specialized economy. In this context, an important issue is to study the relationship of natural resources exports, commodity prices and the exchange rate.

Argentina has historically been a commodity producer and exporter; its exports still remain highly concentrated on a few raw materials and lightly processed primary products. In 2011, seven out of the ten top export items were raw materials: soybean meal, soybeans, soybean oil, corn, wheat, gold and crude oil. These accounted 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 the result of favorable international commodity prices, but of 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 modeling 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 over the last decade.

Despite the fact that Argentina has a reputation of being a resource-rich country largely based on agricultural production, the effect of commodity prices on the economy has not often been examined. 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 income re-distribution took place, there appeared clear symptoms of natural resource curse in the economy, in their view 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 the tax burden rose along with distortionary mechanisms associated with inflation.

Our aim is to empirically study the effects of commodity prices on the Argentine economy. The first link in the chain of effects requires comprehending the effect of prices on both exports and the real exchange rate, which are simultaneously related to each other. To our knowledge this is the first attempt to econometrically study these transmission mechanisms of world commodity prices for Argentina. Most empirical studies have econometrically modelled exports or exchange rates by using single equations. Those focused on the effect of commodity prices on economic

[1] Center of International Economics - Ministry of Foreign Affairs.

growth were based on cross-section or panel data (e.g. Collier and Goderis, 2007; Lederman and Xu, 2009; van der Ploeg and Poelhekke, 2009, among others). Several studies also analyzed the effect of commodity price shocks on the growth of developing economies (Deaton and Miller, 1995; Brown and Yücel, 1999, Izquierdo et al., 2008). In particular, Izquierdo et al. (2008) suggest that one-time increases in commodity prices generate effects on the output level 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 the real exchange rate. The economic structure of Argentina has undergone 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 background and discusses the transmission mechanisms of commodity prices on exports and the real exchange rate. Section 3 describes the data. Section 4 presents the econometric approach for estimating the long-run structure and short-run dynamics. Section 5 discusses the specification of the deterministic components and presents the empirical results. Readers 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 the 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 are the difference between domestic supply and demand of exportable goods, taking 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). 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,

$$x_t = \alpha_0 + \alpha_1 y_t + \alpha_2 p_t^x + \alpha_3 e_t \tag{1}$$

where x stands for the commodity export volume index, y for the production capacity of the exportable sector, 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 and 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 for 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 in an appreciation of the real exchange rate. Therefore, the fact that commodities exports and the real exchange rate may be jointly determined in such cases implies that they should be simultaneously modeled.

Many efforts have been made to empirically model the exchange rate behavior due 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 analyzed economies.

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

Furthermore, an increase in 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 terms of trade shocks which were found to be important empirically, though productivity and government spending differentials continue to be important also.

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

$$e_t = \beta_0 + \beta_1 d_t + \beta_2 yprod_t + \beta_3 p^x_t + \beta_4 x_t \tag{2}$$

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, that is, it appreciates the real exchange rate because of 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. Better commodity terms of trade tend to appreciate the real exchange rate through income or wealth effects. A rise in the commodity sector's productivity may raise the relative price of non-tradable goods (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 favor of such an effect is weaker than commonly believed (Froot and Rogoff, 1995), except for countries with widely disparate levels of income of growth.[2]

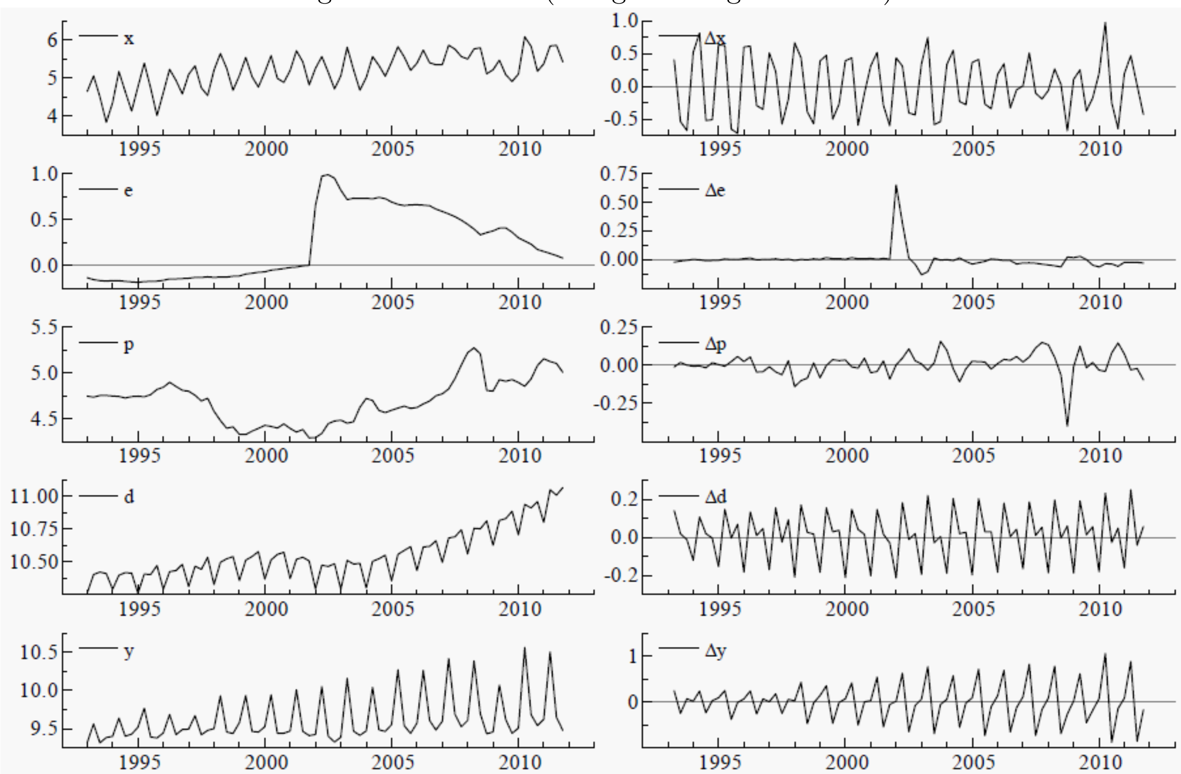
In this paper, we empirically assess the existence of the two long-run (cointegration) relationships expressed in Equation (1) and (2) by applying Johansen's (1996) maximum-likelihood cointegration system approach, as described in the Section 4. This approach allows us to jointly estimate these equations 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 from these equations for the Argentine case, an equilibrium correction representation is also estimated to allow for short-run behavior since we are also interested in studying commodity exports and exchange rate dynamics in a small economy.

[2] In the empirical section we include the agricultural output in the system. 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, shown in Figure 1, are quarterly, over 1993Q1 to 2011Q4 (T=76). The index of commodity prices[3] (p^x) and the real peso/dollar exchange rate (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 peso/dollar exchange rate. The first period (1993-2001) was characterized by the convertibility regime which backed the monetary base with external reserves to guarantee the one peso to one dollar rate of exchange. The economic recession at the end of the convertibility regime (1998-1999) coincided with declining world commodity prices.

The exchange rate regime finally collapsed in January 2002, after the government announced a 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. We can observe that the growth rate of domestic consumption (private and public consumption) has intensified since 2003 when expansionary demand policies also took place. It is worth noting that the world crisis that started 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,

[3] The 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 reflect the product share in Argentine trade.

agricultural production exhibits a changing seasonality which is higher over time, suggesting that the agricultural sector became more dependent on seasonal effects (see Appendix A for seasonally adjusted figures).

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 crises and changes in its economic structure in the last two decades, like different exchange rate regimes as well as fiscal and monetary policy changes. We discuss the deterministic components in the sub-section 5.1 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. In accordance with these statistics and by including a linear broken trend, a I(1) system was estimated to study cointegration, as shown in the next section. Readers not interested in the empirical/econometric methodology can safely jump to Section 6.

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.[4] 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),

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

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.

This 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.[5]

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 indicate which variable adjusts to reach the equilibrium. The coefficients in α give the information about which variables adjust and thereby weak exogenity can be tested by zero restrictions in the respective coefficient, as suggested by Johansen (1992) and Urbain (1992).

[4] However, 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.

[5] Tests of the hypothesis of r cointegration vectors can be based on the trace statistics.

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 a particularly difficult task. The following sections focus on analyzing the appropriate deterministic components to be included in the system.

5 Empirical Results

This section first presents cointegration results obtained 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 the time properties of the series, we introduced the following deterministic components: centered 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 for the whole sample 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 determinist trends (see Juselius, 2006, p.98). In the case we analyzed they may show also determinist trend differences after 2002 when a new economic regime started and many variables grew (or decreased) steadily (see Appendix A). A step dummy 2002Q1-2011Q4 was also incorporated (unrestrictedly) to allow the growth rates to change due to the similarity conditions of the broken trend, as suggested by Nielsen and Rahbek (2000).

After taking account of the 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.[6] After including the determinist components in the VAR, as previously described, normality tests, reported in Table 1, show that multivariate normality is not rejected.[7] 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).

[6] Adding 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-parameterized model. Furthermore, residual autocorrelation in a first tentative VAR(2) model in more often associated with structural misspecification, rather than with left-out dynamic (see Juselius, 2006: p.72).

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

Table 1: 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 their Monte Carlo study. 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 2 reports the estimates for eigenvalues, λ_i and the 95th percentile of the $\Gamma-$distribution when considering a broken trend and a shift dummy in the cointegration relations, $C_{.95}$ (see Nielsen, 1997 and Doornik, 1998).

Table 2: 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 3 reports the 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 3: 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 A1 (see Appendix A) 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 4 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 in order 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). The last two columns report the resulting restricted cointegration relations.

Table 4: 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 centered seasonals, 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 4 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 not being 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 in the soybeans 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 following graphs we can observe the stability of the cointegration results, reported in Table 4, 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. We found some changes in the long-run coefficient estimates during 2002/2003 when the economy began to recover after a deep recession from

1999 to 2001. However, 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).

Figure 2a: Recursive eigenvalues

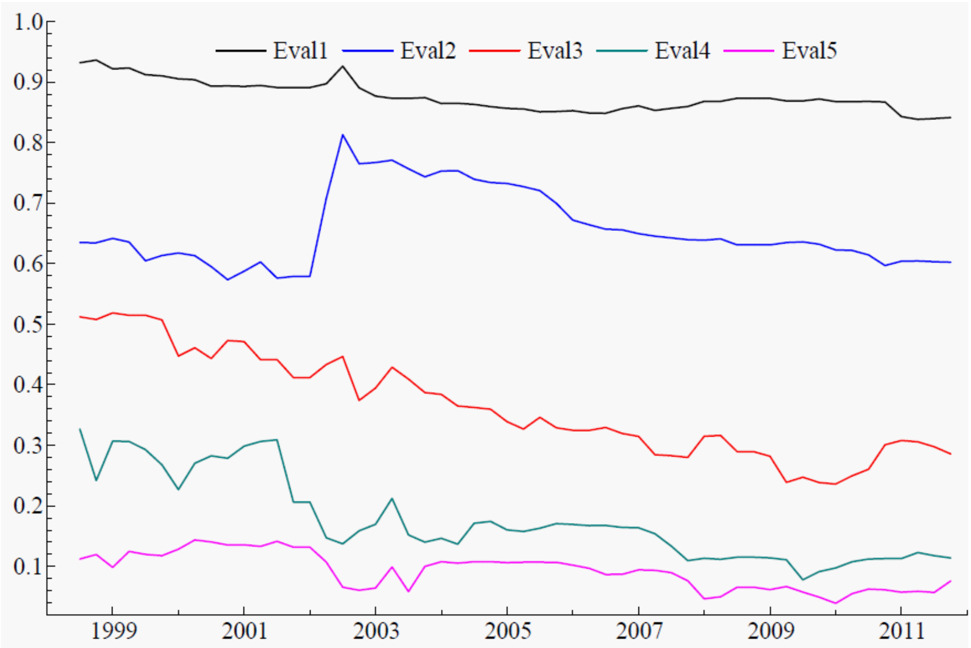
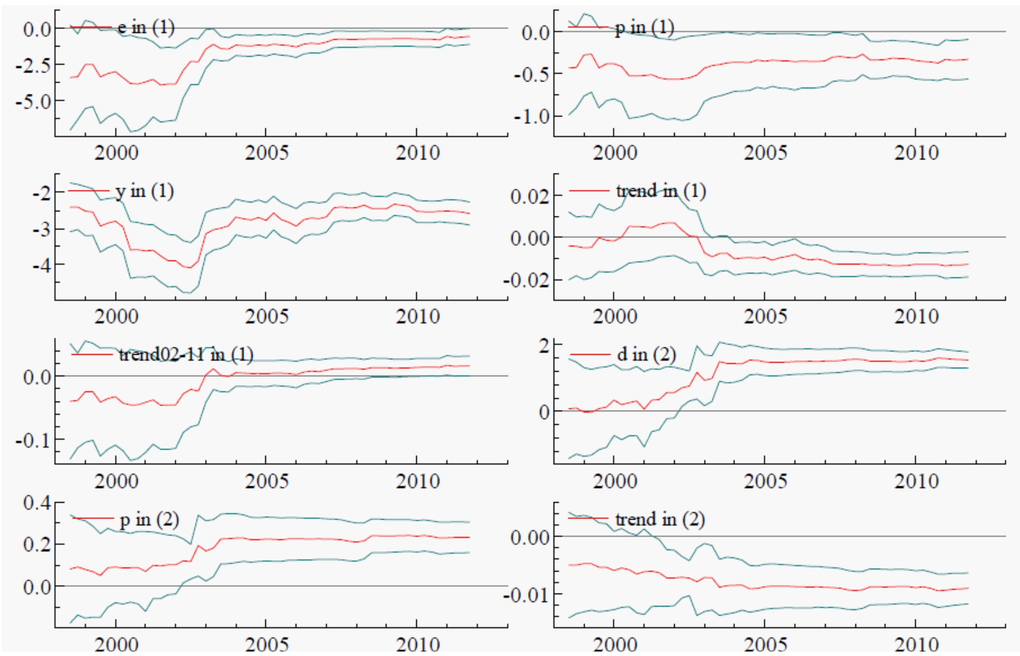


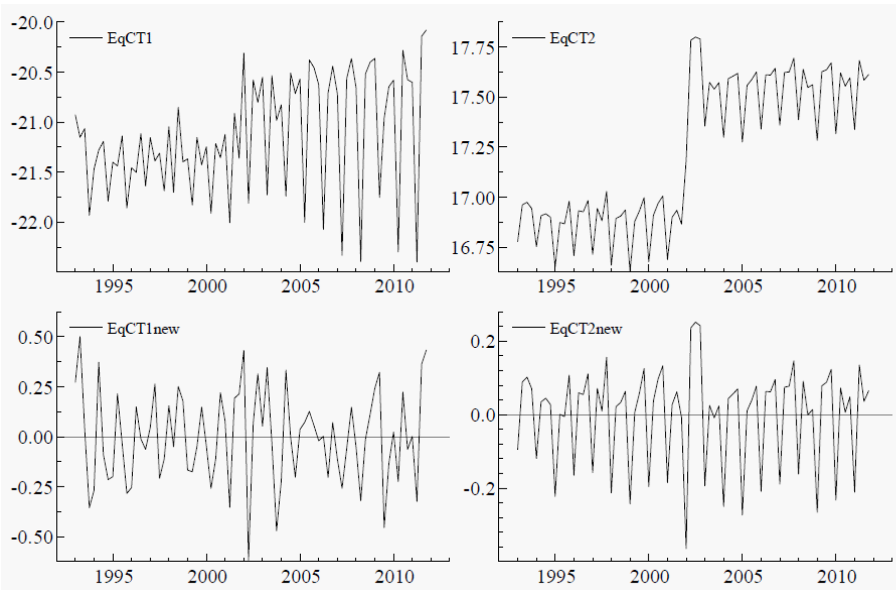
Figure 2b: Recursive eigenvectors (beta $\times \pm 2SE$)



5.2 A simultaneous equilibrium-correction model

Following the proposed normalization of the two identified vectors (as shown in Table 4), we then estimate a vector equilibrium correction model (VEqCM) that considers the simultaneous 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 to adjust the mean of the equilibrium correction terms additionally. Figure 3 illustrates the behavior 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 the agricultural sector have become more dependent on seasonal factors. Therefore, EqCT1 was corrected by obtaining the residuals after regressing it on centered seasonals, broken centered seasonals for 2003-2011 and a step dummy for the period 2002-2011, whereas EqCT2 was 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 using Full Information Maximum Likelihood (FIML). By first removing insignificant variables of the system, based on a Likelihood Ratio test, we arrived at the parsimonious model presented in Table 5. Therefore, a restricted simultaneous EqCM is estimated by FIML. The column heading at the top of the table indicates the dependent variable in each of the model equations, while the row headings indicate the conditioning variables.

Table 5: FIML Estimation of the EqCM

Variable	Δx_t	Δy_t	Δe_t	Δd_t
constant	0.01 (0.02)	0.03 (0.01)	0.001 (0.002)	0.002 (0.003)
$EqCT1new_{t-1}$	-0.23 (0.09)	0.10 (0.04)		
$EqCT2new_{t-1}$			-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^x_t			-0.07 (0.02)	
Δp^x_{t-2}				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: centered seasonal dummies (cseasonals), broken cseasonals 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 5 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 5, while the others are much lower. However, no significant effects were detected for these variables. The growth of commodity prices (Δp) and the depreciation rate (Δe) are the only significant contemporaneous effects.

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 both 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,

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

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

The estimated cointegration relations, as reported in Table 4, show the factors affecting both natural resources exports volume 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 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. An increase in commodity prices will encourage commodity exports.

Exports are more elastic to variations 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 the 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 (4), the differences of growing rates also changed since 2002 as the broken trend ($trend_{02-11}$) is significant too.

From Equation (5) it can be seen that the real exchange rate negatively depends on domestic consumption. An increase in domestic consumption (say 10%) implies an increase in non-tradable good prices pushing down the real exchange rate (in near 15%). 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. However, the elasticity with respecto to domestic demand is higher than the elasticity with respect to commodity prices. Furthermore, no evidence of the 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 exogenous 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 vice versa. 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 export volume and the agricultural production adjust to the deviations from the steady-state, while the real exchange rate, apart from commodity prices, can be considered as given. 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 as presented in section 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) and agricultural production (y) on the one hand, and the real exchange rate (e) and the domestic consumption (d) on the other. 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). Through this estimation, simultaneous short run effects of all variables can also be tested (see Table 5).

We first analyze the estimated commodity exports growth equation,

$$\Delta x_t = \underset{(0.02)}{0.01} - \underset{(0.09)}{0.23}\, EqCT1new_{t-1} - \underset{(0.10)}{0.42}\, \Delta x_{t-2} + \underset{(0.13)}{0.24}\, \Delta y_{t-1} + \underset{(0.12)}{0.31}\, \Delta y_{t-2} + \underset{(0.18)}{0.40}\, \Delta e_{t-4} \tag{6}$$

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.[8]

Commodity exports growth is 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 affects commodity exports. The last year's depreciation rate has also a positive and significant effect on export 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,

$$\Delta y_t = \underset{(0.01)}{0.03} + \underset{(0.04)}{0.10}\, EqCT1new_{t-1} - \underset{(0.08)}{0.53}\, \Delta y_{t-1} - \underset{(0.07)}{0.65}\, \Delta y_{t-2} - \underset{(0.09)}{0.49}\, \Delta y_{t-3} - \underset{(0.30)}{1.15}\, \Delta d_{t-1} \tag{7}$$

The agricultural production growth has also an autoregressive behavior. 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 long-run effect on exports.

The domestic absorption has a short-run negative effect on agricultural production. This finding suggests that, over the sample period, 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,

$$\begin{aligned} \Delta e_t \; = \; & \underset{(0.002)}{0.001} - \underset{(0.05)}{0.16}\, EqCT2new_{t-1} + \underset{(0.03)}{0.42}\, \Delta e_{t-1} - \underset{(0.03)}{0.15}\, \Delta e_{t-2} - \underset{(0.02)}{0.13}\, \Delta e_{t-4} \\ & - \underset{(0.07)}{0.30}\, \Delta d_{t-2} - \underset{(0.08)}{0.37}\, \Delta d_{t-3} - \underset{(0.08)}{0.27}\, \Delta d_{t-4} - \underset{(0.02)}{0.07}\, \Delta p^x_t \end{aligned} \tag{8}$$

[8] After 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.

The depreciation rate is negatively affected by changes in domestic absorption and in world commodity prices. That is, an increase in domestic consumption (via an expansionary fiscal policy or an increasing private consumption) appreciates the real exchange rate. Furthermore, there is a simultaneous negative effect of commodity prices on the 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,

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

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 (or a reduction) of the exchange rate tends to depress (or stimulate) domestic demand. Changes in world commodity prices in the last two quarters have also a significant, but positive effect of domestic absorption. The commodity prices boom encourages expansionary fiscal policies and 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 also been 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 behavior 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 analyze for many developing countries whose economies have undergone deep transformations as a result of the recent commodity boom. For those facing the positive (and unusually long-lasting) shock of high commodity prices, the resulting export growth of natural resources may have also had important consequences derived from the exchange rate appreciation. We develop a case-study for Argentina, a long term and well-known commodity producer and exporter.

By applying Johansen's (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 goods' prices pushing down the real exchange rate reflecting a higher demand for non-tradable goods over their supply during the sample period. 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 interested in studying commodity exports and exchange rate dynamics, an equilibrium correction representation was estimated to allow for short-run behavior.

In a nutshell, the estimated model has shown that commodity prices have been a key variable to explain both exports and the exchange rate. Thus, we econometrically modeled the first chain link of the effects of commodity prices on the Argentine economy. Future research should extend these findings to empirically study the effects on economic growth and income distribution.

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A Analysis of deterministic components

To analyze the degree of persistent behavior in the variables, univariate unit root tests are reported in Table A1. We used 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, commodity prices and agricultural GDP appear to be I(1).

Table A1: 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.

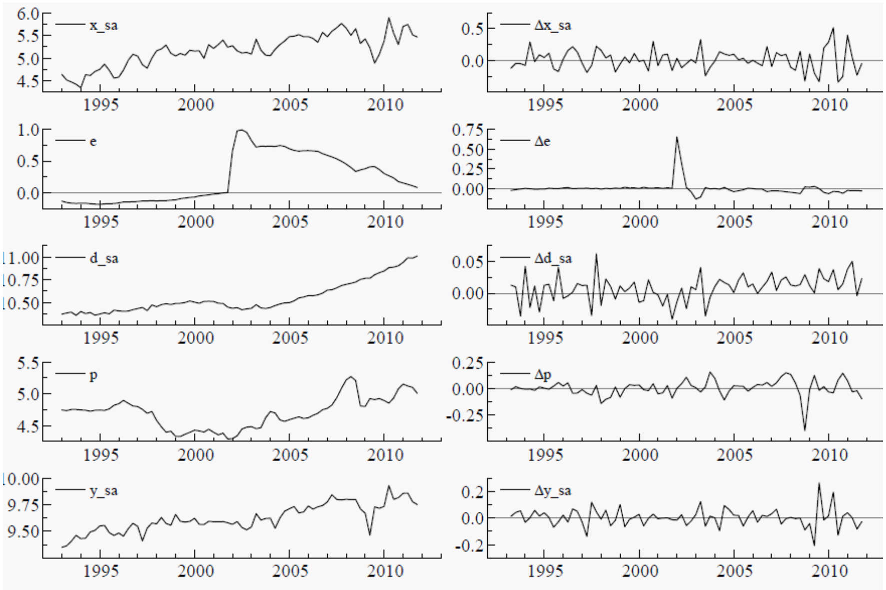
Table A1 also suggests that the domestic consumption is I(2) as the log-difference has a unit root and its second-difference is stationary.. Although this variable would be the only I(2), its effect on a system estimation should be considered. In particular, the typical smooth behavior of a stochastic I(2) trend can often be approximated with an I(1) stochastic trend around a broken linear deterministic trend (see Juselius, 2006: p. 294).

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).[9] From Figure A1 we can observe that the domestic consumption has had a changing growth rate since mid-2000s. During the

[9] Seasonal adjustment was implemented through X12-ARIMA

last decade, fiscal policy became strongly pro-cyclical and the central government has a strongly expansionary fiscal policy.

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

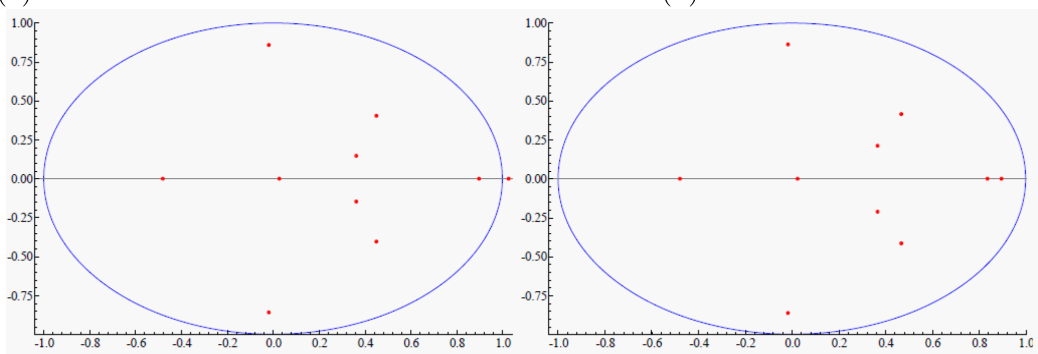


In order to perform an I(1) analysis we allowed for a break in the linear trend since 2002 when the different variables started to grow (or decrease) steadily, but at different rates, as suggested by Figure A1.

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.

Figure A2: Root of the companion matrix

(a) Without a broken linear trend (b) With a broken linear trend



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) behavior 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.