An ex ante analysis of the welfare effects of changes in world prices of agricultural commodities: the case of Uruguay

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Abstract: We analyze the impact on welfare of an increase in commodity prices in Uruguay. Given that this country has a large share of households with low and medium-low income, a positive shock on commodity prices has the potential to hurt a sizable part of the population through the rise in the cost of the consumption basket. However, the pass through on domestic prices would be alleviated by positive changes in labor income that middle-income households experience as the wages also respond to hike of international commodity prices. We find that in Uruguay households at the upper end of the distribution would benefit with the increase of the international prices of agricultural commodities, with low-income households losing as much as 7.5%. In terms of poverty, the increase would be 34%, while increases in indigence would be lower: 19%. Also, the results show that households in a situation of indigence and/or poverty, would move in average further away from the threshold lines, meaning that within each category, poor and indigent households become more homogeneous among them.

JEL Codes: F10, F13, F14, F16, I30.

Keywords: trade, commodity prices, poverty, Uruguay.

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1. Introduction and motivation

There has been an increasing interest on the study of how the deepening of international relations may affect social welfare, employment, inequality and poverty, with the aim of being able to provide policy recommendations to minimize undesirable effects. This new interest has adopted mostly a micro perspective eased by the increasing availability of statistics at the household level, especially for developing and less developed countries.

In this paper we concentrate on the case of Uruguay, which is a country with a large share of its population with low and middle-low income, so there is a need to consider potentially negative effects, such as is the impact on poverty that may follow to a rise in the price of commodities that are used as intermediate inputs in the production of staple goods, which explain a large share of total expenditure in poorer households. As an example of the current importance of this issue, UNCTAD (2013) devoted one chapter of its Commodities and Development Report to the topic of the direct effects of the 2003-2011 commodity boom on poverty and food insecurity.

In the next sections we assess the ex-ante impacts on welfare and poverty at the household level, that can arise due to the increase of the international prices of agricultural commodities. In section 2 we review the existing literature on the nexus between trade policy. In section 3 some descriptive statistics are presented. In Section 4, a simple theoretical model is presented with the aim of serving as the base for our empirical approach. However, it is necessary to say that we do not intend to test structurally the model. Section 5 develops the empirical framework, and presents the results, to estimate the set of parameters that constitute the main ingredient to the analysis carried out in section 6, where we simulate the welfare and poverty effects of a rise of the international prices of agricultural commodities. Section 7 summarizes our main results.

2. The Trade-Poverty nexus: some previous evidence

The economic literature on the links between open trade policies and its assumed positive impact on economic growth and development has reached a consensus when results are measured on average. However, because of the broad set of interrelated factors affecting social welfare outcomes as a result of trade liberalization, when dealing with the potentially beneficial impacts at the level of households, the so called consensus is under dispute. In fact, trade policies have strong redistributive impacts and in most cases it is possible to identify economic groups that benefit and other that are negatively affected. Given the particular importance of local institutional arrangements and market functioning in determining the transmission of border prices to local levels, if poor individuals are among the ones that lose, the long run opportunities for the development of a country or region may be compromised.

McCullloch, et al. (2001) and Winters, et al. (2004) have contributed to deepen and clarify the scope of the debate, summarizing that, the empirical evidence, both in the cases of cross-country and country-case studies, has so far not provided homogeneous results, with liberalization episodes in which the living conditions of the poorer declined. A common feature in terms of the choice of the methodology to assess the direct impact of trade liberalization on poverty is the preference for partial equilibrium techniques instead of general equilibrium (GE) approaches. Indeed, crucial in the choice of the partial equilibrium approach is the possibility of identifying household income and consumption effects. A similar analysis applying GE techniques to quantify distributive effects as a result of price shocks will be limited due to the lack of sufficient disaggregation to fully trace the impact of policies on poverty. Of course, the partial equilibrium approach has the drawback that it

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1 A large body of literature, to which we do not refer to here, has focused almost exclusively on less developed countries, where food security is a very important issue, especially for the poorest households.
leaves aside some second order effects, and even some direct effects. Our exercise is not free of this problem.

The partial equilibrium approach in the existing literature dealing with the trade liberalization poverty nexus starts with the canonical work of Deaton (1989), and gains impetus with the important methodological contribution in Porto (2006). Porto’s methodology allows the identification of two crucial transmission channels: a) the change in relative prices due to a trade reform and b) how these price variations affect households as consumption and income earners. This approach has been eased by the availability of household surveys, especially for developing and less developed countries.

The evidence for Uruguay is mainly due to a study by Borraz, et al. (2012) who analyzes the impact of MERCOSUR on poverty and inequality through two main transmission channels: prices and income. They found that while trade liberalization favored a reduction in poverty indicators, it had an almost zero effect on income inequality, concluding that trade integration policies cannot be regarded as ‘poverty-alleviating’ per se. It could be of interest taking into account the evidence for Argentina, a country that like Uruguay has a large share of the population with medium-low income and with strong comparative advantages in the production of agricultural commodities, so that the effects of an increase in the world commodity prices on welfare could be compared to those of Uruguay.

The Argentine case is studied in Porto (2006 and 2010), Barraud and Calfat (2008), and Barraud (2009), all of which estimate the impact of trade openness on families using household survey data. Barraud and Calfat (2008), Baraud (2009), and Porto (2006) focused on measuring the effects on poverty that resulted from trade liberalization in the nineties. Barraud and Calfat (2008) show that trade liberalization had a pro-poor effect via the reduction in the price of tradable goods and through the effects on the labor market in the sector of non-tradable goods. In the opposite direction, Barraud (2009) obtained that in the case of households related to the manufacturing sector, trade liberalization between 1988 and 1998 would have had a negative impact on poverty. In the pathbreaking methodological work of Porto (2006), the author finds that the implementation of MERCOSUR benefited the average Argentine household across the entire income distribution. As the author points out, the reason behind this result is that Argentine trade policy protected the rich over the poor, prior to the reform, and granted some protection to the poor, after the reform. Finally, Porto (2010) studies the impact of improving access to international agro-manufacture export markets on poverty in Argentina through two channels, the effects caused by price changes on food expenditure and on wages. Porto’s measurement of improved market access is equivalent to an increase in the international price of agro-manufacturing commodities. The main finding is that a better market access would cause poverty to decline in Argentina, result that is mostly driven by relatively large wage elasticities to the prices of agro-manufacture goods (up to 0.85). Recently, Moncarz, et al. (2015) concludes that the increase of the international prices of agricultural commodities during 2002-2011 would have had a negative effect on welfare over the entire income distribution, increasing also the levels of poverty and indigence.

Following Porto (2006) our objective is to contribute to the understanding of how the recent increase in the price of agricultural commodities could have affected welfare and poverty at the household level. We assume an urban model, where households consume products, sell labor but do not produce agricultural commodities. Leaving aside this last effect means that our analysis concentrates on about 93% of the population in Uruguay. With the exception of Moncarz, et al. (2015) for

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2 MERCOSUR is a custom union originally signed by Argentina, Brazil, Paraguay and Uruguay. Venezuela joined recently as the fifth full member, while Bolivia and Chile are associate members under a free trade agreement scheme.

3 As with most of the existing literature, the only effect on income that is taken into account is that on wages. This is, a priori, an important drawback, since an increase in the prices of agricultural commodities should have a positive effect, surely not minor, on the land’s rent.
Argentina, none of the previous evidence for the countries analyzed here has dealt directly with the implications of increasing agricultural commodity prices on the poor.4

3. Agricultural commodity prices, consumer prices and poverty

During the first decade of the current century, international commodity prices increased substantially. The case of agricultural commodities was not the exception. The increases of the prices of the agricultural products exported by Uruguay ranged between 32% and 47% (beef, rice, wheat, and soybeans). In Figure 1.a. we show the evolution of the average prices for the basket of commodities considered, which increased 41% when comparing the averages for 1992-2001 with those for 2002-2011. Figure 1.b. shows the evolution of the main export commodities that comprise the basket. It can be seen that all product shows an upward trend since the 2000's.

It is also important to remark the change in the weight of the main products exported by Uruguay in the total exports during the last years. Table 1.a. reports the weights of the main export products within a group comprising beef, rice, wheat, and soybeans. In 1992 the main export products were the beef (54.2%) and the rice (45.6%). While the participation of the beef remained as being important in the following years, the weight of rice diminished. On the other hand, the exports of wheat and soybeans gained ground in the referred group.

The Figure 1.c. shows the relationship between the price index of agricultural commodities (PWA) and the consumer price index for food and beverages during 1992-2011. Two periods can be distinguished. First during 1992-2000 the prices of commodities exported by Uruguay fell by 30% on average. However, the domestic price index for food and beverages increased by a 579%. Thus, along this time frame, the trend in domestic prices seems to be more related to domestic factors (monetary policy) rather than to the changes in international export prices. It could be suspected that the drivers of consumers prices changed during the following years, as the data report that during 2001-2011 the prices of the commodities exported by Uruguay rose by 52% while the prices of food and beverages increased by 190%.

During the 2000's the external factors gained ground in explaining the domestic prices. The weakness of the dollar triggered a hike in the world commodity prices. For the case of commodity exporters (i.e. Uruguay) it meant an increase in current account balances and a tendency to appreciate the local currency. During 2001-2011 the Central Bank of Uruguay followed an "inflation targeting" scheme in order to control for the creation of money caused by an increase in its exports. This macroeconomic pattern implies a central role of external inflation in explaining domestic prices.

The Figure 1.d. confirms the stated hypothesis about the pass-through of international prices on domestic prices for Uruguay during 2001-2011. It shows the relationship between the consumer price index for food and beverages and the nominal exchange rate during 1992-2000. For the first period analyzed (1992-2000) the nominal exchange rate increased by 389% explaining much of the variation in the price of food and beverages (that rose by 579%). In this first period, the international prices remained low, and the domestic prices are mainly explained by domestic inflation (which in this case is approximated by the evolution of the nominal exchange rate).

In the second time span (2001-2011) the exchange rate increased by 60% (furthermore, the nominal rate decreased by 28% during 2003-2011), while the improvement in the domestic prices more than doubled that figure. To summarize, in this period, the international prices increased and the domestic inflation is thought to be caused by international prices given the Central Bank efforts to control for the inflation. This is the hypothesis to be tested in this paper.

4 de Hoyos and Medvedev (2011) analyze the poverty impact of higher food prices from a global perspective. Lederman and Porto (2015) is a recent review on the effects of changes in commodity and other prices on household welfare, with focus on less developed and developing countries.
The assumption is that the increase of the international prices of agricultural commodities, through a rise in the price of goods that constitute the food basket, which are intensive in the use of these commodities as inputs, has the potential to hurt a sizable proportion of the population in countries where most of households have low and medium-low income. As Table 1.b. shows clearly, Uruguay fall into this last category, with about two-third of households with a level of per capita income lower than the country’s mean. Moreover, a significant proportion of households have an income below half the average income of the country. Due to this pattern of income distribution, it is no surprise that a large share of households spend an important part of their income in food and beverages (see Figure 2), whose prices, as shown in Figure 3, showed an apparently positive relationship with the international prices of agricultural commodities. In the following section we present a model that explains the behavior of domestic prices on the basis of shocks in international prices of commodities.

4. Theoretical framework

In this section we develop a very simple and stylized model to highlight the influence that international commodity prices have on internal prices, both of goods and factors. To keep things as simple as possible, we left aside some issues to which we try to deal with in the empirical application. Because of the lack of appropriate data, we do not intend, in the sections that follow, to do a structural testing of the model we develop in next paragraphs, but it serves the purpose of justifying the empirical approach we apply.

The theoretical framework assumes a small open economy that produces and trades S primary commodities, of which \( S_s \subset S \) are agricultural commodities. Assuming the number of primary commodities is at least as large as the number of factors, then factor rewards are fully determined by commodity prices:

\[
W = p(P^u_s)
\]

(1)

where \( W \) is the vector of factor rewards, and \( P^u_s \) is the vector of commodity prices in local currency.

Since our economy is small, we have:

\[
P^o_s = EP_s (1 + T)
\]

(2)

where \( E \) is the nominal exchange rate, \( P^o_s \) is the vector of international commodity prices, and \( T \) is the vector that reflects the ad-valorem equivalent of the country trade policy, so we obtain:

\[
W = p(P^o_s, E, T)
\]

(3)

There are also \( M \) traded manufacturing sectors, of which \( M_r \subset M \) produce food goods. Following what has become standard in the trade literature, the \( M \) manufacturing sectors are monopolistically competitive with production exhibiting increasing returns to scale (IRS). In each m sector each producer, domestic or foreign, produces a differentiated variety using all factors of production and primary commodities. There are also \( N \) non-traded sectors that are also monopolistically competitive, with each domestic producer producing a differentiated variety under IRS using only the production factors.\(^6\)

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\(^5\) One of this issues is the relationship between international commodity prices and the exchange rate.

\(^6\) The assumption that non traded goods are produced only using production factors, is only a simplification which has no impact on the relationships of international prices of commodities with either consumer prices nor with factor rewards.
Assuming also that production factors are perfectly mobile across all sectors, the price, in local currency, of each domestic variety of the M and N sectors can be expressed as a function of international commodity prices, and other parameters such that nominal exchange rate, domestic taxes/subsidies, trade policy etc.

To be more specific, let us assume that there are two primary commodities, A1 and A2, whose internal prices are given by:

\[ p_{A1}^d = E \cdot p_{A1}^* (1 + \tau_{A1}) \]  
\[ p_{A2}^d = E \cdot p_{A2}^* (1 + \tau_{A2}) \]  

where \( E \) is the nominal exchange rate, \( \tau_{A1} \) and \( \tau_{A2} \) are the ad-valorem equivalents of the country trade policy on goods A1 and A2 respectively, and the superscript * makes reference to international values. Then, given the small country assumption we get:

\[ w_1 = f_1(P^*, T, E) \]  
\[ w_2 = f_2(P^*, T, E) \]

with \( P^* = (p_{A1}^*, p_{A2}^*) \) and \( T = (\tau_{A1}, \tau_{A2}) \).

Each variety \( i \) produced by the manufacturing sector \( m \) is produced under IRS using the two factors of productions and the two primary commodities, with total costs equal to:

\[ TC_{i,m} = C_{i,m} \left( \alpha_m + \beta_m x_{i,m} \right) \]

where \( \alpha_m \) is the fixed input requirement, \( \beta_m \) is the input per unit of output produced by each firm, \( x_{i,m} \), and \( C_{i,m} \) is a Cobb-Douglas composite defined as:

\[ C_{i,m} = w_1^{\alpha_m} w_2^{\beta_m} \left( p_{A1}^d \right)^{\tau_{A1}} \left( p_{A2}^d \right)^{1-\tau_{A1}-\tau_{A2}} \]

Each industry is monopolistically competitive, with each firm in sector \( m \) facing a constant elasticity of demand equal to \( \sigma_m \), so the producer price of a domestically produced variety \( i \) in sector \( m \) is given by:

\[ p_{i,m} = C_{i,m} \beta_m \left( \frac{\sigma_m}{\sigma_m - 1} \right) \]

Then, the consumer price, and under the simplifying assumption that there are no domestic taxes or subsidies, is:

\[ p_{c,m}^i = p_{i,m} \]  
\[ \sigma_m \]

For an imported variety, and defining \( \tau_{i,m}^{imp} \) as the ad-valorem equivalent of trade costs on imports, the consumer price is equal to:

\[ p_{c,m}^{imp} = E \cdot p_{i,m}^* (1 + \tau_{i,m}^{imp}) \]

Finally, assuming that in each sector all firms are identical, and a CES function that determines the consumption of each variety of sector \( m \), we have that the consumer price index for all varieties (produced domestically and imported) of a given sector \( m \) is:

\[ \sigma_m \]

---

7 The constant elasticity of demand follows from the assumption that the consumption of each variety produced by sector \( m \) is the result of a Constant Elasticity of Substitution (CES) function.
\[ P_m = \left[ N_m \left( p_{m}^{*} \right)^{1-\sigma_m} + N_m^* \left( p_{m}^{*} \right)^{1-\sigma_m} \right]^{1/(1-\sigma_m)} \]  

where \( N_m \) and \( N_m^* \) are, respectively, the number of varieties produced domestically and abroad.

Working in a similar way as for the M sectors we obtain the following relationships for each non-traded sector \( n \):

\[ TC_{i,n} = C_{i,n} \left( \alpha_n + \beta_n x_{i,n} \right) \]  

\[ C_{i,n} = w_{1}^{n} w_{2}^{i,n} \]  

\[ p_{i,n} = C_{i,n} \beta_n \left( \frac{\sigma_n}{\sigma_n - 1} \right) \]  

\[ p_i' = p_{i,n} \]  

\[ P_n = \left[ N_n \left( p_{i,n}^{*} \right)^{1-\sigma_n} \right]^{1/(1-\sigma_n)} \]  

Using (5) and (7)-(10) into (11), and (5) and (13)-(15) into (16), it emerges clearly that consumer price indices for the M and N sectors are only functions of international commodity prices (in the case of imported varieties, the effects of international commodity prices enter indirectly through their effect on the producer prices of such varieties) and the parameters of the model. These relationships, as well as the effects of international commodity prices on factor prices, are the ones we look to estimate in the next section.

5. Computing long-run elasticities of consumer prices and wages with respect to world prices

Most of the existing literature on the subject relies on performing an impulse-response analysis to compute the pass-through of international prices to internal ones. For example, Furlong and Ingenito (1996), Krichene (2008), Zoli (2009), Ferrucci, et al. (2010), Rigobon (2010), and Ianchovichina, et al. (2014), among others fit a Vector Autorregresive (VAR) model and then estimate the corresponding response of internal prices to a given shock in international commodity prices. However, this approach fails to provide an "standard" measure of elasticity: that is, rather than providing the percentage change of a determined internal price to a one-percentage change in the international price (i.e. the elasticity of the internal price with respect to the international price), that "VAR approach" captures the response of the internal price to a "shock" to the international price, with this shock usually defined as one standard deviation.

In our case, instead, we estimate the long-run elasticities by identifying a Vector Error Correction (VEC) model. This allows us to obtain the elasticities according to the usual definition. Additionally, the identification of the cointegrating relationships implies adding theoretical assumptions, which provides an economic content to the analysis of the long run dynamics of the price time series.

Let’s consider the VEC representation of a VAR of order \( l \), given by:

\[ \Delta p_t = \Pi \Delta p_{t-1} + \Gamma_1 \Delta p_{t-1} + \ldots + \Gamma_{l-1} \Delta p_{t-l+1} \]  

Given that \( p_t \) is a Kx1 vector that contains at least one I(1) variable, \( \Pi \) is a singular KxK matrix with rank equal to \( r \). Further, \( \Pi \) can be written as \( \Pi = \alpha \beta' \), where the KxK matrix \( \beta \) is the cointegrating matrix. We are interested in analyzing the rx1 vector \( ec_{t-1} = \beta' p_{t-1} \), that contains the cointegration

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8. Anderson and Tyers (1992) is an example of the use of a an error-correction model to compute elasticities for changes in border prices relative to domestic producer prices.
relations between prices. In particular, if the variables included in $p_t$ are expressed in logarithms, the coefficients in $\beta$ represent the elasticities that measure the response of consumer prices to the international ones. Providing that the cointegrating rank is known, the reduced-rank maximum likelihood estimator $(\alpha_e, \beta_e)$ is available; however it only estimates consistently the cointegrating space. Therefore, it is necessary to identify K-r variables utilizing prior information. We assume that the first part of $\beta$ is an identity matrix, so it takes de form $\beta'=[I_r, \beta'_{k-r}]$, where $I_r$ is an identity matrix of order $r$, while $\beta'_{k-r}$ is an $r \times (K-r)$ matrix with the coefficients to be identified. For identification purposes, our assumption is that consumer prices are driven by the international ones.

As pointed out by Juselius (2006), all the long-run effects are captured in the $\Pi$ matrix, and the VECM provides a way to distinguish between long- versus short-run coefficients. Given that $\beta_{t-1}^p$ represents a stationary linear combination of the variables, the coefficients of $\beta$ describe the relations of these variables in the steady-state. When these relationships are interpreted according to the economic theory, the concept of cointegration matches the notion of long run equilibrium.

In this section we compute long-run elasticities of consumer prices and salaries. To do this, we perform a cointegration analysis between variables. Particularly, we consider two different alternatives to identify these coefficients. In what follows we make a brief description of them.

5.1. Identifying long-run elasticities by estimating a Vector Error Correction Model (VECM) for each response variable

In this case the strategy is to estimate a different VEC model for each response variable. When the objective is to estimate the consumer price elasticities, each model will provide the elasticity of a given consumer price index with respect to the international price index of agricultural commodities and to the nominal exchange rate. To do this, we define the following four vectors:

$$p_{fb}^t=(pc_{fb}^t, e_t, p_{wa}^t)'$$
$$p_{clo}^t=(pc_{clo}^t, e_t, p_{wa}^t)'$$
$$p_{equ}^t=(pc_{equ}^t, e_t, p_{wa}^t)'$$
$$p_{oth}^t=(pc_{oth}^t, e_t, p_{wa}^t)'$$

(18)

In each model, $p_{j}^t$ is a 3x1 vector that contains (the log of) the consumer price index for a given category of goods ($j=fb, clo, equ, oth$), (the log of) the exchange rate index ($e$), and (the log of) the international price index of agricultural commodities ($pwa$). In the four vectors, the first element corresponds to a consumer price index (i.e. food and beverages, clothing, equipment, and other goods), the second one represents the nominal exchange rate, while the remaining element is given by international price index of agricultural commodities, which is supposed to led the long run trend in the referred consumer price.

For the vector $ec_{t-1}=\beta_{i}^t p_{t-2}$, if rank($\beta_{i}^t$)=1, then $\beta_{i}^t=(1, \beta_{i11}^t, \beta_{i21}^t)$ and the cointegration equation is represented as:

$$pc_{j}^t=\beta_{j11}^t e_{t-1}+\beta_{j21}^t pwa_{t-1}$$

(19)

The proposed estimation framework is also capable to test additional hypothesis in an open economy. For instance, given the order the variables enter the vector $p_{j}^t$, if two cointegration relations were found, then two hypotheses could be considered: first, consumer prices are driven by international commodity prices, and second, the exchange rate depends also on international commodity prices. It is supposed that the cointegration rank ranges between 1 and 2 because at least one of the three variables in the vector is I(1).

The estimation strategy is as follows. Unit root tests are applied on each variable separately to determine the order of integration. The following stage entails working separately with the vectors $p_{fb}^t, p_{clo}^t, p_{equ}^t, p_{oth}^t$. The optimal lag length for the VAR representation of each vector is computed...
according different criteria. Cointegration tests are run to determine the cointegrating rank associated to $\Pi^{10}$, $\Pi^{10}$, $\Pi^{10}$, and $\Pi^{10}$. Finally, each VECM is estimated separately for each vector (after imposing identifying restrictions) to obtain $\beta_f$, $\beta_c$, $\beta_{eu}$ and $\beta_{oth}$, which are thought to contain the long run elasticities of consumer prices ($pcfb$, $pcclo$, $pcequ$, $pcoth$) with respect to the exchange rate ($e$) and the international prices of agricultural commodities ($pwa$).

When the objective is to estimate the wage elasticities, each model will provide the elasticity of a given wage index with respect to the exchange rate and the international price index of agricultural commodities. Thus, we define the following three vectors:

\[
\begin{align*}
              w_1 &= (w_{1t}, e_t, pwa_t)' \\
    w_2 &= (w_{2t}, e_t, pwa_t)' \\
         w_3 &= (w_{3t}, e_t, pwa_t)'
\end{align*}
\]  

In each model, $w_j$ is a 3x1 vector that contains (the log of) the average hourly wage for workers with a certain level of formal education ($j = 1, 2, 3$), (the log of) the exchange rate index ($e$) and (the log of) the international price index of agricultural commodities ($pwa$). For the vector $ec_{t-1} = \beta' w_{j,t}$, if $\text{rank}(\beta)=1$, then $\beta' = (1, \beta_{11}, \beta_{21})$ and the cointegration equation is represented as:

\[
w_{jt-1} = \beta_{11} e_{t-1} + \beta_{21} pwa_{t-1}
\]  

and we follow the estimation strategy as it was depicted above.

### 5.2. Identifying long-run elasticities by performing a unique cointegration analysis for the entire set of the response variables

In this case the strategy is to estimate a single VECM including all variables. For the case of consumer prices, and defining the following vector:

\[
p_{t} = (pcfb_t, pcclo_t, pcequ_t, pcoth_t, e_t, pwa_t)'
\]  

for the cointegration equations $ec_{t-1} = \beta' p_{t-1}$, if $\text{rank}(\beta)=4$, then:

\[
\begin{pmatrix}
1 & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 0 & 1 & 0 \\
0 & 0 & 0 & 1
\end{pmatrix}
\]

where, for example, the first cointegration equation is given by:

\[
\begin{align*}
    pcfb_{t-1} &= \beta_{51} e_{t-1} + \beta_{61} pwa_{t-1} \\
\end{align*}
\]  

In equation (7), $\beta_{51}$ and $\beta_{61}$ represent the elasticities of $pcfb$ with respect to $e$ and $pwa$, respectively. The estimation strategy is as follows. The optimal lag length for the VAR representation of the vector $p_t$ is computed according different criteria. Cointegration tests are run to determine the cointegrating rank associated to $\Pi$. Finally, the VECM is estimated (after imposing identifying restrictions) to obtain $\beta$, which are thought to contain the long run elasticities of consumer prices ($pcfb$, $pcclo$, $pcequ$, $pcoth$) with respect to the exchange rate ($e$) and the international prices of agricultural commodities ($pwa$). To estimate the pass-through of international commodities prices to wages, we define:

\[
w_{t} = (w_{1t}, w_{2t}, w_{3t}, e_t, pwa_t)'
\]  

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9 The results of these preliminary analyses are available upon request.

10 Unskilled ($j=1$), Semi-skilled ($j=2$) and Skilled ($j=3$). Individuals are divided into the three groups depending on their formal education: incomplete high school or less ($j=1$), complete high school or incomplete tertiary/university ($j=2$), complete tertiary/university ($j=3$).
For the cointegration equations $e_{ct-1} = \beta'w_{t-1}$, if rank($\beta$) = 3, then:

$$\beta' = \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \beta_{41} \\ \beta_{42} \\ \beta_{43} \end{pmatrix} \begin{pmatrix} \beta_{51} \\ \beta_{52} \\ \beta_{53} \end{pmatrix}$$

where, for example, the first cointegration equation is given by:

$$w_{1t-1} = \beta_{41} e_{t-1} + \beta_{51} pwa_{t-1}$$

(25)

such that $\beta_{41}$ and $\beta_{51}$ represent the elasticities of $w_1$ with respect to $e$ and $pwa$, respectively. Then, the estimation strategy follows the one previously described.

Next we present the results obtained for Uruguay.

5.3 Results

In Table 2 we present the elasticities of consumer prices for the period 2002-2011, using monthly data. The coefficients were estimated jointly by adding all the consumer prices in the same vector in order to perform the VEC analysis. The cointegration test did not reject the hypothesis of 4 cointegration relationships. Thus, the gathered evidence suggests that the long-run behavior of each consumer price is linked to the international prices through a specific stochastic trend, which is distinct to that of the remaining consumer prices. The elasticities of $pcf_b$, $pcclo$ and $pcequ$ are significantly different from zero at a 1% level, while the coefficient for $pcoth$ is not significantly different from zero (probably because the VEC system finds difficult to capture the dynamics of a variable that consists of an aggregate of the prices of different category of goods).

Finally, to capture the wage dynamics in Uruguay, we estimated separately 3 VECMs; in each case the dependent variable is the corresponding wage index associated to a different level of education (see Table 3) using monthly data for the period 2002-2011. The first thing to remark is that in every case the hypothesis of one cointegration relation was not rejected; this would imply that there is no evidence favoring the hypothesis that the exchange rate was driven by the international prices of agricultural commodities (this relationship would emerge in a second cointegration relation for each VECM). Second, the estimated coefficients are all significantly different from zero (1% level for the three cases). Third, the response of wages to the international prices of agricultural commodities seems to be independent of the education level. All in all, the coefficients are quite high, and show the vulnerability of workers to the international business cycles. Overall, the estimated elasticities for consumer prices are in line with what can be expected a priori. The elasticity of food and beverage is considerably higher than for the other three category of goods.

6. Effects on welfare and poverty of an increase of the international prices of agricultural commodities

Our primary goal is to simulate the effects on welfare and poverty that may follow an increase of the international prices of agricultural commodities. Once we have obtained the elasticities of consumer prices and wages with respect to the international prices of agricultural commodities, we can simulate the welfare effects that would follow to a given shock in the latter.

In particular, the welfare effect on household $h$ will be measured by the negative of the compensating variation relative to its initial expenditure:

$$\frac{d x_h}{e_h} = - \left( \sum_{j \in N,M} s_j^h e_j a_{j} \right) d \ln p_{it} + \left( \sum_j q_j^h e_j a_{j} \right) d \ln p_{it}$$

(26)
where $s^g_h$ is the budget share spent on varieties produced by sector $g$, $\psi_{g,s}$ is the elasticity with respect the international price index of agricultural commodities $\left(p_{s}\right)$ of the consumer price index for goods of sector $g$, $\theta^h_j$ is the salaried labor income of member $j$ as a share of total income of household $h$, and $\varepsilon_{g,s,p}$ is the wage elasticity that captures the proportional change in the wage rate of household member $j$ as a response to the change of the international prices of agricultural commodities. The first term in the RHS of (26) is the welfare effect that takes place through consumption, while the second term measures the effect through changes in labor income. Considering the way in which equation (26) is computed, a negative value means a welfare loss, while a positive value means a welfare gain. In equation (26) we do not consider second-order effects that take place through changes in consumption patterns in response to changes in consumer prices, neither changes in the supply of labor. Also, and because of the lack of appropriate data, we do not take into account the effects on non-labor income, neither those due to the consumption of own-produced goods. Finally, as already mentioned above, all the following analysis ignores any impact on the rural population.

6.1. Welfare effects

Using the elasticities obtained in Section 5, and budget shares from the national surveys of household expenditures, and assuming a 100% increase of the international prices of agricultural commodities, applying equation (26) we obtain the effect on welfare for each household. Then, we run non-parametric regressions of the welfare effects as a function of household per capita expenditure.

Before looking at the results of the simulations, in Table 4 we report some descriptive statistics about the consumption patterns and the sources of income, since these together with the elasticities of the previous sections determine the magnitude of the simulated effects. As it emerges clearly, in the case of the expenditure shares, the category of food and beverages has a much greater importance for poorer households. These differences help to explain why in our simulations the poorer households are more negatively affected through the consumption effect. On the other hand, the differences between households are less important when we look at the sources of income, however it still is possible appreciate the lower participation of salaried labor for the poorest households. It is important to highlight a drawback of our the analysis, due to the fact that we cannot account for the effects working through other sources of income, which most likely would help to diminish the magnitudes of the aggregate negative effects we obtain.

In the simulation of the welfare effects we need to deal with two sources of randomness. The first comes from the sampling variability of expenditure shares and of the participation of salaried labor in the household income; while the second source emerges because of the error associated to the estimation of the responses of consumer prices and wages to international agricultural commodity prices. To jointly account for these sources of variability, we work as follows. The randomness due to sampling variability of households (and therefore of budget shares and that of the participation of salaried labor in household incomes) is controlled by weighting each observation by the inverse probability of its inclusion into the sample. To deal with the variance of the estimated elasticities, we follow Porto (2006) and resample from their empirical asymptotic distribution. From de VEC models in section 5, we obtain, for each category of consumption goods and for each wage index, an estimate $\hat{\beta}_j$ of the elasticity with respect to the international price index of agricultural

\[^{11}\text{We also followed Porto (2006) and took random samples of households from the expenditure and income surveys. The results were almost identical, but our approach demanded much less computation time. Sample weights come from Uruguay's official statistic office.}\]
commodities, and also an estimated standard error $\hat{\sigma}_j$. Under standard assumptions $\hat{\beta}_j \sim N(\beta_j, \sigma_j)$, where $\beta_j$ and $\sigma_j$ are the true parameters values. Then, in each loop a new elasticity is assigned to the formula to calculate the welfare effect. The non-parametric regression is run for each of the 200 replications to deliver new estimates of the average welfare effect. After the 200 replications, we compute the standard error of the estimated regression functions to build the 90% confidence bands.

The welfare effect through the consumption of food and beverages is, as expected, negative for all households. As shown in Figure 4, the welfare loss is 15%, showing also great variability across households. In the case of the consumption of non-food and beverages (Figure 5), losses are of much lower magnitude. The main reason behind this result is the lower elasticities than the ones obtained for food and beverages.

Once we add the effects taking place through the consumption of all kind of goods, and because of the larger impact that works through the consumption of food and beverages, the qualitative results resemble those reported in Figure 4, but now with losses ranging between 8% and 18% (see Figure 6).

To obtain the labor income effects, we use the wage elasticities reported in Table 3. Then, using the share of salaried income of each member in the household total income, and once again assuming a 100% increase in the price of agricultural commodities, we calculate the welfare effect coming through changes in wages.

As it is shown in Figure 7, there is a positive effect through the increase in labor income. Middle income households are the one that benefit the most. It can be explained because of the larger share of salaried income in total household income for those in the middle of the distribution. Another possible explanation is the higher elasticity of wages for semi-skilled workers.

Finally, once we add the effects that work through consumption and labor income, we obtain that poorest households are the most affected (see Figure 8). Households at the upper end of the distribution would benefit with the increase of the international prices of agricultural commodities, with low-income households losing up to 7.5%. Uruguay shows high variability between households.

All in all, the distribution of losses along the per capita expenditure of households is, a priori, in line with what could be expected: an increase in the price of agricultural commodities hurting more to poorer households due to the higher weight of food and beverages into household consumption, which are goods intensive in the use of agricultural commodities.

6.2. Effects on indigence and poverty

To grasp an approximate idea of how important is the impact on poverty of an increase of the international prices of agricultural commodities, in Table 5 we report the indigence and poverty rates that would follow after a 100% increase, as well as two additional measures: the gap and severity of indigence and poverty.\textsuperscript{12} Indigence and poverty are measured in absolute terms, comparing the

\textsuperscript{12} In all Figures, the solid line is the average effect, while dashed lines are the 90% confidence bands.

\textsuperscript{13} The rate, gap and severity of indigence and poverty are measured following Foster \textit{et al.} (1984), using the following formula: $R = \frac{1}{N} \sum_{h=1}^{N} \left( \frac{y_h - z_h}{z_h} \right)^{\alpha} I(y_h < z_h)$, where $N$ is the total number of households, $z_h$ is the indigence/poverty threshold for household $h$ (these thresholds are household-specific, depending on the structure of the household in terms of the age and gender of its members), $y_h$ is total income of household $h$, and $I(y_h < z_h)$ is a latent variable equal to 1 if $y_h < z_h$. When $\alpha = 0$ we obtain the rates of indigence/poverty, if $\alpha = 1$ we have the indigence/poverty gap, and when $\alpha = 2$ we have the indigence/poverty severity.
income of the household with the minimum expenditures required not to fall into either of the two categories. The indigence line is defined for a individual of reference\textsuperscript{14}, measuring the minimum expenditure necessary to acquire the basic food basket (CBA), which is calculated to guarantee the intake of a certain number of calories. On the other hand, the poverty line is obtained by multiplying the CBA by the Engel's coefficient, which gives us the total basic basket (CBT). For Uruguay the Engel's coefficient is about 4.6 for the city of Montevideo and 3.25 for the remaining urban areas. Finally, multiplying the CBA and the CBT by the household size\textsuperscript{15}, the indigence and poverty baskets are calculated for each household.

To obtain the new value of the CBA for each household, we update the original indigence line for the time each household was surveyed, considering only the effect that works through the increase in the consumer price of food and beverages.\textsuperscript{16} Then, the new poverty line is obtained using the Engel's coefficient for the time each household was surveyed. The new household incomes are calculated taking into account only the effect on labor income of salaried household members.\textsuperscript{17}

As reported in Table 5, we obtain that a 100% increase of the international prices of agricultural commodities would led to an increase in indigence of 3.2 pp. in Uruguay. For poverty, the increases would be 9.8 pp. However, while in absolute terms poverty rates increases more than those of indigence, the opposite is true when the changes are measured relative to the initial values; with indigence increasing 40%. For instance, poverty rate would increase 34% of their initial values, while in the case of indigence the changes would be 49%.

If instead of using a headcount measure, we look at the deepness of indigence and poverty, we obtain that in relative terms, the gap and severity of both indices increase more than their corresponding rates. These results mean that there is not only an increase in indigence and poverty in response to the raise of the international prices of agricultural commodities, but also that those households who were already indigent and/or poor, as well as those who become indigent and/or poor, move in average further away from the threshold lines. These results mean also that within each category, poor and indigent households become more homogeneous among them.

7. Conclusions

With a large share of households with low and medium-low income, the increase in agricultural commodities prices has the potential to hurt a sizable part of the population through a rise in the price of goods that explain an important share of households expenditures, those that constitute the food-basket. The ex-ante simulations show that we can expect this to be the case. A less obvious channel works through changes in factor incomes. In the case of salaried labor income, the magnitude of the positive effect is not large enough to compensate for the losses because of the changes in the prices of consumption goods, with the exception of the richest households in Uruguay.

\textsuperscript{14} The indigence line, in per capita terms, is built on the basis of the food requirements of a group of representative households according to their socioeconomic and demographic characteristics.

\textsuperscript{15} There is the assumption that there exist economies of scale within the household, so the household size is lower than the number of individuals.

\textsuperscript{16} It would have been more appropriate to work with the changes in the prices of goods that constitute the CBAs, data we do not have access to, however, the correlation of the consumer price index for food and beverages with the indigence and poverty lines is around 0.99 in both cases.

\textsuperscript{17} In this point is necessary to highlight an official figures of indigence and poverty in Uruguay include as part of the household income the amount corresponding to the imputed rental value for those households that own the house they live in. In all our analysis, and with the objective of making the results more comparable with those for other Latin-American countries, we exclude the imputed rental value. Working this way means that rates of indigence and poverty as calculated here are higher than those reported officially.
In terms of poverty, there would be an increase of 34%, while for indigence the changes would be even greater, around 49%. Also, the results show that households in a situation of indigence and/or poverty, would move in average further away from the threshold lines, with households within each category, poor and indigent, becoming more homogeneous among them. Finally, our results highlight the need for well designed compensatory measures to help those that could be most negatively affected. Finally, the analysis should be interpreted into the context of its limitations, mostly because of the impossibility to account for the effects working through changes in factor rewards other than salaried labor, especially the land rent, which in countries like Uruguay benefited greatly from the increase in the prices of agricultural commodity.

Also, an important element that would have been interesting to account for, is that of non-salaried labor income, which can be of importance for small entrepreneurs as well as independent workers and small shopkeepers.
### Table 1.a.
Weights of the main products exported by Uruguay in total exports.

<table>
<thead>
<tr>
<th>Year</th>
<th>Beef</th>
<th>Wheat</th>
<th>Rice</th>
<th>Soyabeans</th>
</tr>
</thead>
<tbody>
<tr>
<td>1992 (*)</td>
<td>0.542</td>
<td>0.000</td>
<td>0.456</td>
<td>0.003</td>
</tr>
<tr>
<td>1993 (*)</td>
<td>0.542</td>
<td>0.000</td>
<td>0.456</td>
<td>0.003</td>
</tr>
<tr>
<td>1994</td>
<td>0.542</td>
<td>0.000</td>
<td>0.456</td>
<td>0.003</td>
</tr>
<tr>
<td>1995</td>
<td>0.580</td>
<td>0.031</td>
<td>0.389</td>
<td>0.000</td>
</tr>
<tr>
<td>1996</td>
<td>0.464</td>
<td>0.022</td>
<td>0.514</td>
<td>0.000</td>
</tr>
<tr>
<td>1997</td>
<td>0.457</td>
<td>0.066</td>
<td>0.477</td>
<td>0.000</td>
</tr>
<tr>
<td>1998</td>
<td>0.458</td>
<td>0.034</td>
<td>0.508</td>
<td>0.000</td>
</tr>
<tr>
<td>1999</td>
<td>0.514</td>
<td>0.039</td>
<td>0.447</td>
<td>0.000</td>
</tr>
<tr>
<td>2000</td>
<td>0.607</td>
<td>0.004</td>
<td>0.390</td>
<td>0.000</td>
</tr>
<tr>
<td>2001</td>
<td>0.471</td>
<td>0.000</td>
<td>0.524</td>
<td>0.005</td>
</tr>
<tr>
<td>2002</td>
<td>0.536</td>
<td>0.000</td>
<td>0.433</td>
<td>0.031</td>
</tr>
<tr>
<td>2003</td>
<td>0.525</td>
<td>0.000</td>
<td>0.398</td>
<td>0.077</td>
</tr>
<tr>
<td>2004</td>
<td>0.631</td>
<td>0.003</td>
<td>0.244</td>
<td>0.122</td>
</tr>
<tr>
<td>2005</td>
<td>0.647</td>
<td>0.011</td>
<td>0.228</td>
<td>0.114</td>
</tr>
<tr>
<td>2006</td>
<td>0.652</td>
<td>0.015</td>
<td>0.204</td>
<td>0.129</td>
</tr>
<tr>
<td>2007</td>
<td>0.537</td>
<td>0.019</td>
<td>0.254</td>
<td>0.190</td>
</tr>
<tr>
<td>2008</td>
<td>0.510</td>
<td>0.071</td>
<td>0.241</td>
<td>0.178</td>
</tr>
<tr>
<td>2009</td>
<td>0.411</td>
<td>0.130</td>
<td>0.230</td>
<td>0.228</td>
</tr>
<tr>
<td>2010</td>
<td>0.409</td>
<td>0.133</td>
<td>0.162</td>
<td>0.296</td>
</tr>
<tr>
<td>2011</td>
<td>0.417</td>
<td>0.114</td>
<td>0.174</td>
<td>0.296</td>
</tr>
</tbody>
</table>

Note: Import weighted average of prices for beef, rice, wheat, and soybeans. Source: own based on WITS and www.indexmundi.com (retrieved on November 12, 2012).

### Table 1.b.
Proportion of households with per capita income lower than value of reference

<table>
<thead>
<tr>
<th>Year</th>
<th>1/4 of medium income</th>
<th>1/2 of medium income</th>
<th>Medium income</th>
</tr>
</thead>
<tbody>
<tr>
<td>2003</td>
<td>9.3</td>
<td>32.7</td>
<td>67.5</td>
</tr>
<tr>
<td>2011</td>
<td>4.8</td>
<td>25.0</td>
<td>64.1</td>
</tr>
</tbody>
</table>

Source: own calculations based on Encuesta Continua de Hogares (Uruguay). In all cases, household income excludes the implicit rent for those households who own the house they live in.
Table 2
Consumer price elasticities $\beta p_{t-1}$

<table>
<thead>
<tr>
<th>Cointegration equation</th>
<th>$pcfb_{t-1}$</th>
<th>$pcclo_{t-1}$</th>
<th>$pcequ_{t-1}$</th>
<th>$pcoth_{t-1}$</th>
<th>$e_{t-1}$</th>
<th>$pwa_{t-1}$</th>
<th>TREND</th>
</tr>
</thead>
<tbody>
<tr>
<td>$ec_{1,1}$</td>
<td>1.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>-0.446</td>
<td>-0.513</td>
<td>-0.005</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.029)</td>
<td>(0.0369)</td>
<td>(0.000)</td>
<td></td>
</tr>
<tr>
<td>$ec_{2,1}$</td>
<td>0.000</td>
<td>1.000</td>
<td>0.000</td>
<td>0.000</td>
<td>-0.754</td>
<td>-0.169</td>
<td>-0.004</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.052)</td>
<td>(0.064)</td>
<td>(0.001)</td>
<td></td>
</tr>
<tr>
<td>$ec_{3,1}$</td>
<td>0.000</td>
<td>0.000</td>
<td>1.000</td>
<td>0.000</td>
<td>-0.615</td>
<td>-0.331</td>
<td>-0.005</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.062)</td>
<td>(0.076)</td>
<td>(0.001)</td>
<td></td>
</tr>
<tr>
<td>$ec_{4,1}$</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>1.000</td>
<td>-0.852</td>
<td>-0.027</td>
<td>-0.008</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.049)</td>
<td>(0.060)</td>
<td>(0.001)</td>
<td></td>
</tr>
</tbody>
</table>

[Std. Dev.] (p - Value) [t - Value] *** p<0.01, ** p<0.05, * p<0.1.

The Akaike Criteria indicated that the optimal VAR lag length is equal to 9. The cointegration test was run using Lütkepohl and Saikkonen (L&S) procedure. The null hypothesis $H_0$: rank($\beta$)=4 cannot be rejected, so that the VECM was specified assuming that the cointegration rank is equal to 4. Remaining VECM's specification details are as follows: deterministic variables: TREND; endogenous lags (in differences): 8; sample range: [2002 M10, 2011 M12]; T=111; estimation procedure: One stage. Johansen approach. Further estimation details are available upon request.

Source: Own calculations.
### Table 3
Wage elasticities

*Coefficients of cointegration relations $\beta^j_w t-1$*

<table>
<thead>
<tr>
<th>Cointegration equation</th>
<th>$w_{j,t-1}$ (#)</th>
<th>$e_{t-1}$</th>
<th>$pwa_{t-1}$</th>
<th>TREDN</th>
</tr>
</thead>
<tbody>
<tr>
<td>$w_t^1 = (w_{1t-1}, e_{t-1}, pwa_{t-1})$</td>
<td>1.000</td>
<td>-0.513</td>
<td>-0.367</td>
<td>-0.009</td>
</tr>
<tr>
<td>(0.000)</td>
<td>(0.122)</td>
<td>(0.142)</td>
<td>(0.002)</td>
<td></td>
</tr>
<tr>
<td>{0.000}</td>
<td>{0.000}***</td>
<td>{0.010}**</td>
<td>{0.000}***</td>
<td></td>
</tr>
<tr>
<td>[0.000]</td>
<td>[-4.208]</td>
<td>[-2.579]</td>
<td>[-5.392]</td>
<td></td>
</tr>
<tr>
<td>$w_t^2 = (w_{2t-1}, e_{t-1}, pwa_{t-1})$</td>
<td>1.000</td>
<td>-0.506</td>
<td>-0.401</td>
<td>-0.007</td>
</tr>
<tr>
<td>(0.000)</td>
<td>(0.125)</td>
<td>(0.146)</td>
<td>(0.002)</td>
<td></td>
</tr>
<tr>
<td>{0.000}</td>
<td>{0.000}***</td>
<td>{0.006}***</td>
<td>{0.000}***</td>
<td></td>
</tr>
<tr>
<td>[0.000]</td>
<td>[-4.058]</td>
<td>[-2.752]</td>
<td>[-4.061]</td>
<td></td>
</tr>
<tr>
<td>$w_t^3 = (w_{3t-1}, e_{t-1}, pwa_{t-1})$</td>
<td>1.000</td>
<td>-0.500</td>
<td>-0.406</td>
<td>-0.007</td>
</tr>
<tr>
<td>(0.000)</td>
<td>(0.118)</td>
<td>(0.137)</td>
<td>(0.002)</td>
<td></td>
</tr>
<tr>
<td>{0.000}</td>
<td>{0.000}***</td>
<td>{0.003}***</td>
<td>{0.000}***</td>
<td></td>
</tr>
<tr>
<td>[0.000]</td>
<td>[-4.251]</td>
<td>[-2.953]</td>
<td>[-4.102]</td>
<td></td>
</tr>
</tbody>
</table>

(Std. Dev.) (p - Value) [t - Value] *** p<0.01, ** p<0.05, * p<0.1.

For the first two VECM the Akaike, Hannan-Quinn and Schwarz Criteria indicated that the optimal VAR lag length is equal to 2. For the third VECM the Hannan-Quinn and Schwarz Criteria indicated that the optimal VAR lag length is equal to 2, and the Akaike indicated that the VAR lag length is equal to 3.

The cointegration tests were run using Lütkepohl and Saikonen (L&S) procedure. In all cases the null hypothesis $H_0: \text{rank}(\beta^j) = 1$ cannot be rejected, so that the VEC were specified assuming that the cointegration rank is equal to 1. Remaining VECM's specification details are as follows:

deterministic variables: TREDN; endogenous lags (in differences): 1; sample range: [2002 M2, 2011 M12]; T=118; estimation procedure: Two-stage Johansen approach. Further estimation details are available upon request.

Source: Own calculations.
<table>
<thead>
<tr>
<th>Quintile</th>
<th>FB</th>
<th>CLO</th>
<th>EQU</th>
<th>OTH</th>
<th>Salaried Labor</th>
<th>Labor</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>38.5</td>
<td>5.1</td>
<td>4.3</td>
<td>52.1</td>
<td>40.5</td>
<td>58.1</td>
</tr>
<tr>
<td>2</td>
<td>31.6</td>
<td>3.9</td>
<td>3.6</td>
<td>60.9</td>
<td>47.0</td>
<td>62.2</td>
</tr>
<tr>
<td>3</td>
<td>26.1</td>
<td>4.0</td>
<td>3.7</td>
<td>66.3</td>
<td>46.4</td>
<td>61.6</td>
</tr>
<tr>
<td>4</td>
<td>22.9</td>
<td>3.8</td>
<td>4.0</td>
<td>69.4</td>
<td>46.8</td>
<td>61.3</td>
</tr>
<tr>
<td>5</td>
<td>15.8</td>
<td>3.8</td>
<td>5.5</td>
<td>74.9</td>
<td>39.2</td>
<td>57.4</td>
</tr>
</tbody>
</table>

source: own calculations based on Encuesta Nacional de Gastos e Ingresos de los Hogares 2005-2006. FB: food and beverages; CLO: clothing; EQU: equipment; OTH: other goods and services. Labor: includes self employment. (#) Shares in consumption as a percentage of total household expenditure; sources of income as a percentage of total household income (excluding the imputed rental value for households which own the houses they live in).
### Table 5
Indigence and Poverty rates in urban areas
Pre and post a 100% increase in international prices of agricultural commodities

<table>
<thead>
<tr>
<th>Variable</th>
<th>Uruguay</th>
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<tbody>
<tr>
<td></td>
<td>Value</td>
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<tr>
<td>Rate</td>
<td>Pre</td>
</tr>
<tr>
<td></td>
<td>Post</td>
</tr>
<tr>
<td>Indigence Gap</td>
<td>Pre</td>
</tr>
<tr>
<td></td>
<td>Post</td>
</tr>
<tr>
<td>Severity</td>
<td>Pre</td>
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<tr>
<td></td>
<td>Post</td>
</tr>
<tr>
<td>Rate</td>
<td>Pre</td>
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<tr>
<td></td>
<td>Post</td>
</tr>
<tr>
<td>Indigence Gap</td>
<td>Pre</td>
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<tr>
<td></td>
<td>Post</td>
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<tr>
<td>Severity</td>
<td>Pre</td>
</tr>
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<td></td>
<td>Post</td>
</tr>
</tbody>
</table>

The rate, gap and severity of indigence and poverty are measured following Foster et al. (1984):

\[
R = \frac{1}{N} \sum_{h=1}^{N} \left( \frac{z_h - y_h}{z_h} \right)^{\alpha} I \left( y_h < z_h \right),
\]

where \( N \) is the total number of households, \( z_h \) is the indigence/poverty threshold for household \( h \) (these thresholds are household-specific, depending on the structure of the household in terms of the age and gender of its members), \( y_h \) is total income of household \( h \), and \( I \left( y_h < z_h \right) \) is a latent variable equal to 1 if \( y_h < z_h \). When \( \alpha = 0 \) we obtain the rates of indigence/poverty, if \( \alpha = 1 \) we have the indigence/poverty gap, and when \( \alpha = 2 \) we have the indigence/poverty severity.

(*) Bootstrapped standard errors.

Note: Total household income excludes the imputed rental value for households which own the houses they live in.

Source: own calculations.
Figure 1.a.
Average international prices of agricultural commodities

Note: Import weighted average of prices for beef, rice, wheat, and soybeans. Source: own based on WITS and www.indexmundi.com (retrieved on November 12, 2012).

Figure 1.b.
Prices of the main products exported by Uruguay. 1992-2011. Monthly data

Source: own based on WITS and www.indexmundi.com (retrieved on November 12, 2012).
Figure 1.c.
Relationship between the evolution of commodity and consumption indexes in Uruguay.

Figure 1.d.
Relationship between the nominal exchange rate and the price index of food and beverages

Source: own based on WITS and www.indexmundi.com (retrieved on November 12, 2012) and Instituto Nacional de Estadística.
Figure 2
Share of expenditure in Food and Beverages

Figure 3
Consumer and agricultural commodity prices

Source: own calculations based on Instituto Nacional de Estadística, www.indexmundi.com (retrieved on November 12, 2012), and WITS of World Bank.
Figure 4
Consumption effect of a 100% increase of the international prices of agricultural commodities Food and Beverages
Figure 5
Consumption effect of a 100% increase of the international prices of agricultural commodities Non-Food and Beverages

[Diagram showing the consumption effect of a 100% increase of international prices for agricultural commodities Non-Food and Beverages, with the x-axis labeled as ln(Hou. Exp. pc) and the y-axis ranging from -0.8 to 0.02.]
Figure 6
Aggregate consumption effect of a 100% increase of the international prices of agricultural commodities
Figure 7
Labor income effect of a 100% increase of the international prices of agricultural commodities
**Figure 8**

Aggregate effect of a 100% increase of the international prices of agricultural commodities
References


Appendix

Data description and sources

**Agricultural commodity prices**: weighted average of the country main agricultural exports (Beef, Wheat, Rice and Soybeans). Exports values are used as weights. Price indices are from www.indexmundi.com (retrieved on November 12, 2012) while exports are from WITS of World Bank.

**Consumer Prices**: are from Instituto Nacional de Estadística.

**Wages**: average wages are from Instituto Nacional de Estadística using the Encuesta Continua de Hogares.

**Nominal Exchange Rate**: is from Banco Central del Uruguay.

**Household expenditures and incomes**: are from Instituto Nacional de Estadística, Encuesta Nacional de Gastos e Ingresos de los Hogares 2005-2006.

### Regional coverage

<table>
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<tr>
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<tbody>
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<td>Consumer prices</td>
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<td>Wages</td>
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<tr>
<td>Welfare / Poverty</td>
<td>Urban areas</td>
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