Poverty Impacts of Changes in the Price of Agricultural Commodities: Recent Evidence from Argentina

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Poverty impacts of changes in the price of agricultural commodities: recent evidence for Argentina (*an ex-ante analysis*)[¥]

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Abstract: Argentina, like other land abundant countries, benefited greatly from the recent increase in the prices of agricultural commodities. However, and in despite of the benefits at the macro level, with a large share of the population with low and medium-low incomes, the increase in agricultural commodity prices has the potential to hurt an important part of the population through a raise in the price of goods that explain a large share of households expenditures, especially those that constitute the food-basket. The ex-ante evidence shows that this is expected to be the case. A less obvious channel, through changes in factor incomes would be more beneficial to the middle income households. The elimination of the VAT on the consumption of food and beverages would be enough to compensate for most of the negative effects derived from the increase in the world prices of agricultural commodities.

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I. Introduction and motivation

Among the current research agenda of international trade there is 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 looking to minimize undesirable effects. This new interest has adopted mostly a micro perspective eased by the increasing availability of statistics at the household level, specially for developing and less developed countries.

Due to the increasing integration in world trade markets, Argentina, like another land abundant countries, has benefited greatly from the recent increase in the prices of agricultural commodities that took place during the last decade. For instance, for the main agricultural products exported by the country (soybeans, soybean meal, soybean oil, sunflower oil, maize and wheat), average prices in the 2002-2012 period have increased between 42% and 84% compared to the average of the preceding ten years. As it is shown in Figure 1, this increase in world prices of agricultural commodities has been part of a more general tendency which also happened in other commodity markets.

To have a clearer look of the importance of these prices changes for Argentina's exports, in Table 1 we decompose the change in export values between the change in prices and the change in quantities. As shown in the Table, during the period 1992-2002 the price index of exports fell by a 9%, while quantities increased by 130%. On the other hand, in the period 2002-2012 instead, the increase in the value of exports was mainly driven by the change in prices with a 100% rise, while quantities increased only 58%. This change in the source of growth is also present, and even at a greater extent in the cases of agricultural commodities and manufacturers intensive in their use, and even more for fuel and energy.[1]

Despite of the benefits, at the macro level, that followed the increase in the price of agricultural commodities, such as the important increase in exports which helped to easy the external restriction that in the past has conditioned the country long-run growth possibilities[2], there is a need to consider other effects that may be less desirable. One of this effects 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 food goods, which explain a large share of total expenditure in poorer households (see Figure 2). As an example of the current importance of this issue, recently, UNCTAD (2013) has 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.

At the peak of the 2002 economic crisis, when the local currency had already depreciated by almost 300%, the rise in the price of agricultural commodities contributed further to the increase in domestic prices, especially those of tradable goods. As an example, in the period

[1] In the last decade, exports of fuel have been subject to important restrictions and heavy taxes, which explains that despite of a 350% price increase, quantities fell by 69%

[2] During the period 1992-2001, Argentina exported by 215.95 billions USD, during the following ten years it did by 510.83 billions. Imports, on the other hand, were 215.91 and 361.68 billions USD respectively. The increase of exports acquires a greater importance when we take into account that since 2002 the country has been almost completely excluded from international financial markets.

2002-2011[3], while the overall consumer price index increased by 420%, that of food and beverage increased by 638.7%, the highest among the nine consumption categories, and only matched by the 612.6% increase in clothing. In Figure 3 we can observe an apparent positive relationship between consumer prices and world prices of agricultural commodities.

In the next sections we assess the ex-ante impact on poverty at the household level, that can arise due to the increase in the prices of agricultural commodities. Also, we evaluate the ex-ante effect that would follow the elimination of the value added tax (VAT) on the consumption of food and beverages.

II. Previous evidence for Argentina

The theoretical developments in trade theory and the policies to foster international trade show that in most cases is possible to identify economic groups that benefit and other that are negatively affected. Given particular institutional arrangements and market functioning, if poor individuals are among the ones that lose, the long run opportunities for the development of a country or region may be compromised.

As Winter et al. (2004) summarizes, 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. As for the methodology implemented, the progress by the recent studies is reflected in going beyond the traditional theoretical postulates of the Stolper-Samuelson theorem, trying to estimate these effects at the household level. This approach has been eased by the availability of household surveys, specially for developing and less developed countries. The Argentine case is studied in Porto (2006 and 2010), Calfat and Barraud (2008) and Barraud (2009), all of which estimate the impact of trade openness on families using household survey data.

The evidence for Argentina (Barraud and Calfat, 2008 and Porto 2006 and 2010) has 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. Porto (2006) finds that the implementation of MERCOSUR[4] 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. Porto (2010) studies the impact of improving

[3] Since 2007 there has been a growing distrust about price statistics carried out by the National Institute of Statistics and Census (INDEC), thus for the period 2007-2011 prices indices are obtained using inflation data calculated by the Government of the Province of San Luis. More recently the distrust has extended to other statistics, such as measurements of poverty, employment, and growth.

[4] 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.

access to international agro-manufacture markets on poverty in Argentina through two channels, the effects caused by prices changes on food expenditure and on wages. The main finding is that a better market access would cause poverty to decline in Argentina.

In the present study, the objective is to contribute to the understanding of how the recent increase in the price of agricultural commodities, which is expected to persist over the medium-run, can affect poverty in Argentina, as well as discuss possible policy responses that would serve to minimize possible undesirable effects. None of the previous evidence for Argentina has dealt with this topic.[5]

III. Theoretical framework

The theoretical framework assumes a small open economy that produces and trades S primary commodities, of which $S_A \subset S$ 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:

$$W = p\left(P_S^D\right)$$

where W is the vector of factor rewards, and P_S^D is the vector of commodity prices in local currency.

Since our economy is small, we have:

$$P_S^D = EP_S^*\left(1+\mathrm{T}\right)$$

where E is the nominal exchange rate, P_S^* is the vector of world commodity prices, and T is the vector that reflects the ad-valorem equivalent of the country trade policy, so we obtain:

$$W = p\left(P_S^*, E, \mathrm{T}\right)$$

There are also M traded manufacturing sectors, of which $M_F \subset M$ produce food goods. The M manufacturing sectors are monopolistically competitive. In each M sector each producer, domestic or foreign, produces a differentiated variety. Manufactures are produced under increasing returns to scale (IRS), 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.

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 world commodity prices, and other parameters such that nominal exchange rate, domestic taxes/subsidies, trade policy etc.

[5] de Hoyos and Medvedev (2011) analyze the poverty impact of higher food prices from a global perspective.

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

$$p_{A1}^{d} = E\, p_{A1}^{*}\left(1+\tau_{A1}\right)$$

$$p_{A2}^{d} = E\, p_{A2}^{*}\left(1+\tau_{A2}\right)$$

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

$$w_1 = f_1\left(P^{*}, \mathrm{T}, E\right)$$

$$w_2 = f_2\left(P^{*}, \mathrm{T}, E\right)$$

where $P^{*} = \left(p_{A1}^{*}, p_{A2}^{*}\right)$ and $\mathrm{T} = \left(\tau_{A1}, \tau_{A2}\right)$.

Each variety *i* produced by the manufacturing sector m is produced under IRS using the two factor 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 α_m is the fixed input requirement, β_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:

$$C_{i,m} = w_1^{\mu_m} w_2^{\delta_m} \left(p_{A1}^{d}\right)^{\gamma_m} \left(p_{A2}^{d}\right)^{1-\mu_m-\delta_m-\gamma_m}$$

Each industry is monopolistically competitive, with each firm in sector m facing a constant elasticity of demand equal to σ_m[6], 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:

$$p_{i,m}^{c} = p_{i,m}$$

[6] 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.

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

$$p_{i,m^*}^c = E\, p_{i,m^*}\left(1+\tau_m^{imp}\right)$$

Finally, assuming that in each sector all firms are symmetric, and the CES function that determines the consumption of each variety of sector m, we have that the price index for all varieties (domestic and imported) of sector m is given by:

$$P_m = \left[N_m \left(p_{i,m}^c \right)^{1-\sigma_m} + N_{m^*} \left(p_{i,m^*}^c \right)^{1-\sigma_m} \right]^{\frac{1}{1-\sigma_m}}$$

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:

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

$$C_{i,n} = w_1^{\eta_n} w_2^{1-\eta_n}$$

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

$$p_{i,n}^c = p_{i,n}$$

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

As it emerges clearly from the price indices for the M and N sectors, they are function, among other factors, of international commodity prices. These relationships, as well as the effect on factor prices, are the ones we need to estimate in the empirical section.

IV. Empirical framework

The methodology will follow that of Deaton (1989) and Benjamin and Deaton (1993), which consists of estimating two links, one that connects world commodity prices to domestic prices (goods and factors), and a second one connecting domestic prices to household welfare.

IV.1 Price elasticities

Most of the existing literature on the subject relies on performing an impulse-response analysis to compute the pass-through of international prices to domestic ones. For example, Furlong and Ingenito (1996), Krichene (2008), Zoli (2009) and Rigobon (2010) among others fit a Vector Autorregresive (VAR) model and then estimate the corresponding response of domestic 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 domestic price to a one-percentage change in the international price (i.e. the elasticity of the domestic price with respect to the international price), that "VAR approach" captures the response of the domestic 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.

Before presenting in the next section the results, we now briefly present the structure of the model we work with. Given a set of K time series variables representing the prices of goods, $p_t=(p_{1t},p_{2t},...,p_{Kt})'$, the dynamic interactions between these variables can be captured through a Vector Autoregressive Model of order l, VAR(p), given by:

$$p_t = A_1 p_{t-1} + ... + A_l p_{t-l} + u_t$$

where A_j (j=1,...,l) is a KxK matrix, and u_t is assumed to represent a white noise process, with time-invariant, positive definite covariance matrix. If the process has a unit root, some or all variables are said to be integrated. Then, we are interested in analyzing the cointegration relationships that appear explicitly in the VEC representation of the previous VAR process:

$$\Delta p_t = \Pi p_{t-1} + \Gamma_1 \Delta p_{t-1} + ... + \Gamma_{l-1} \Delta p_{t-l+1} + u_t \qquad (1)$$

If the VAR process has unit roots, the KxK matrix Π is singular. Assuming that p_t can be at most I(1), it turns out that Δp_t does not contains stochastic trends. As a consequence, the term Πp_{t-1} must also be I(0), being the only one that includes I(1) variables. This term specifies the cointegrating relationships. In particular, the number of cointegrating relationships is given by the rank of Π:

$$rk(\Pi)=r$$

where r is the cointegrating rank. Π can be written as $\Pi=\alpha\beta'$. The Kxr matrix β is called the cointegrating matrix, as the rx1 vector $ec_{t-1}=\beta' p_{t-1}$ contains the cointegration relations between prices. Note that $rk(\alpha)=rk(\beta)=r$. α is known as the Kxr loading matrix, that contains the weights attached to the cointegrating relations in the individuals equation of the model. Of course, matrices α and β are no unique if r<K. Therefore, it is necessary to gather non-

sample information (associated to the economic theory) to fully identify the cointegrating relationships.

If the cointegrating rank is known, the reduced-rank maximum likelihood estimator (α_e,β_e) is available, which only estimates consistently the cointegrating space. However, to estimate α and β consistently it is necessary to add identifying (uniqueness) restrictions (given that Π is not singular, it is necessary to identify K-r variables utilizing prior information). The most widespread practice in the literature is to assume that the first part of β is an identity matrix, so it takes de form $\beta'=[I_r: \beta'_{k-r}]$, where I_r is and identity matrix of order r, while β'_{k-r} is an rx(K-r) matrix with the coefficients to be identified.[7]

For identification purposes, our assumption is that domestic prices are driven by the international ones. Then, let us define $p_t=(p_{1t},p_{2t},...,p_{Kt})'=(pdfb_t, pdclo_t, pdequ_t, pdoth_t, pwa_t)'$; where the first four elements are the (log of) domestic prices of food and beverages (*pdfb*), clothing (*pdclo*), equipment (*pdequ*) and other goods[8] (*pdoth*); while *pwa* is the (log of) the international price of agricultural commodities. This ordering implies that innovations in international prices came first.

The estimation strategy is as follows. Firstly, unit root tests are applied on each variable separately to determine the order of integration. Secondly, the optimal VAR lag is computed according different criteria. Thirdly, cointegration tests are run to determine the cointegrating rank among the group of selected variables. Finally, the VEC model is estimated (after imposing indentifying restrictions) to obtain the cointegrating matrix β. This matrix is thought to contain the long run elasticities of domestic prices (*pdfb*, *pdclo*, *pdequ*, *pdoth*) with respect to the international price of agricultural commodities (*pwa*).

IV.2 Wage elasticities

For the relationship between labor income and the international price of agricultural commodities, we follow what is standard in the literature and estimate an extended Mincer-wage equation with the following general specification:

$$\ln w_{j,t} = \alpha + \sum_{edu=1,2,3} \beta_{edu} \left(d_{edu} \times \ln P^*_{s_A,t} \right) + \sum_{edu=1,2,3} \delta_{s,edu} \left(d_{edu} \times \ln E_t \right) + \mathbf{Z_j} \Pi + u_{j,t} \qquad (2)$$

where $w_{j,t}$ is the log of the average hourly wage for an individual *j*, d_{edu} (edu=1,2,3) are three dummy variables to distinguish between three different levels of formal education: incomplete high school or less (edu=1), complete high school or incomplete tertiary/university (edu=2) and complete tertiary/university (edu=3), $P^*_{s_A,t}$ is an index of world prices of Argentina's main agricultural commodities, E_t is the nominal exchange rate

[7] See Lütkepohl and Krätzig (2004) for more details.

[8] Housing, transport and communication, education, leisure, and other goods and services.

between the local currency and the US dollar, and **Z** is a set of additional explanatory variables.

More specifically, we estimate five alternative specifications of an extended Mincer equation, the five specifications differ from each other depending on the set of controls included in **Z**:

Model 1: age and age squared; dummy variables for males, head of household, not-single status, education (incomplete tertiary/university and complete tertiary/university education - the reference group is incomplete secondary or less), formal job, firm size (6 to 50, and more than 50 employees - the reference group is 1 to 5 employees), type of firm (private sector, and other sector - the reference group is public sector), sector of activity (20 dummies - the reference sector is agriculture, hunting, forestry and fishing), place of residence (31 dummies - the reference city is Buenos Aires), year (16 dummies - the reference year is 1995), quarter (3 dummies - the reference quarter is January-March)

Model 2: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with a linear trend.

Model 3: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with a linear and quadratic trends.

Model 4: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with two linear trends, the first starting from the beginning of the period, and the second one starting from 2004.

Model 5: same as Model 1 plus the interaction of all variables (except for the year and quarter dummies) with a dummy variable equal to 1 from 2004 to 2011.

All models include a constant term.

The interaction of the variables included in the set Z with the time trends or the dummy variable for the period 2004-2011, allows for changes in wages to be explained by other variables that only the price of agricultural commodities and the exchange rate, avoiding the introduction of a bias in the calculation of the wage elasticities.

IV.3 Welfare effects

For the second link, the welfare effect on household *h* will be measured by the negative of the compensating variation relative to total initial expenditure:

$$\frac{dx_0^h}{e^h} = \left(-\sum_{g \in N,M} s_g^h \frac{\partial \ln P_g}{\partial \ln p_{s_A}} + \sum_j \theta_j^h \varepsilon_{w,p_{s_A}}^j \right) d \ln p_{s_A} \qquad (3)$$

where s_g^h is the budget share spent on varieties produced by sector *g*, θ_j^h is the labor income of member *j* as a share of total income of household *h*, $\varepsilon_{w,p_{s_A}}^j$ is the wage elasticity that captures the proportional change in the wage of household member *j* as a response to the change in the world price of an agricultural commodity p_{s_A}; P_g is price index for sector *g*. The way we compute equation (3) a negative value means a welfare loss, while a positive value means a welfare gain.

In equation (3) we do not consider second-order effects that take place through changes in consumption patterns in response to changes in domestic prices. However, if we compare expenditure shares between the survey for 2004/2005 (the one we use for our simulations) and the immediate before for 1996/1997, there are important differences in expenditure shares for some goods, especially in the case of food and beverages (see Figure 2), the one we expect to be the main driving force behind the changes in welfare. A similar result emerges when looking at the sources of income, with an important reduction of labor income, at the expense of transfers, for those at the low end of the expenditure distribution (see Figure 4). While the first change would help to reduce the negative effect on those at the lower end of the expenditure distribution, the second one plays the opposite role since now any benefit deriving from an increase in the retribution to labor has a lower impact.

To include the second-order effects that work through changes in the patterns of consumption in response to changes in domestic prices, we calculate the compensating variation allowing for these responses, more specifically we calculate:

$$\frac{dx_0^h}{e^h} = \left(-\sum_{g\in N,M} s_g^h \frac{\partial \ln P_g}{\partial \ln p_{s_A}} + \sum_j \theta_j^h \varepsilon_{w,p_{s_A}}^j \right) d\ln p_{s_A} - \left(\frac{1}{2} \sum_{g\in N,M} \sum_{k\in N,M} s_g^h \sigma_{gk} \left(\frac{\partial \ln P_g}{\partial \ln p_{s_A}} d\ln p_{s_A} \right) \left(\frac{\partial \ln P_k}{\partial \ln p_{s_A}} d\ln p_{s_A} \right) \right) \quad (3')$$

where σ_{gk} is the compensated price elasticity between goods of sector *g* with respect to goods of sector *k*.[9] In (3') we do not allow for second-order effects that work through changes in labor income.

Both in (3) and (3'), and due to data availability we do not take into account the effects on non-labor income. Also, because of data restrictions, we assume households do not produce for their own consumption.

Finally, once the welfare effects have been recovered, we run non-parametric regressions of the changes in welfare as a function of household expenditure per capita.

[9] Price elasticities are from Florensa and Moncarz (2014).

V. Results

V.1 Elasticities of domestic prices and wages

For equation (1), we use monthly data covering the period 1992-2011. To take into account the possibility of a structural change in the data generating process, we split the whole period into two sub periods; firstly, the VEC model is adjusted to resemble the dynamics of the prices during the so-called convertibility period (1992-2000), in which the exchange rate remained fixed by law. Secondly, we fit the model utilizing the post-convertibility period during which time there was exchange rate flexibility and a loose monetary policy (2003-2011). Years 2001 and 2002 were left aside in order to discard the effect of a period of exceptional macroeconomic disturbances that could affect the measurement of the elasticity coefficients. There is another factor that justifies the division of the data into two periods. As shown in Figure 1, the series representing the international price index of agricultural commodities displays a deterministic upward trend since the beginning of the 2000's. All estimations were performed using the J-Multi sofware (Lütkepohl and Krätzig; 2004).

For the first sub-period (1992-2000), the optimal lag length (in levels) is 2. The Johansen-trace test suggested that the cointegration rank (using seasonal dummies (D) and an intercept) equals 3. The estimated cointegrating relationships are:

$$ec_{1,t} = pdfb_t + \underset{(0,827)}{0,171}\, pdoth_t - \underset{(-2,823)***}{0,137}\, pwa_t - \underset{(-5,933)***}{5,011} + D_{1,t}$$

$$ec_{2,t} = pdclo_t - \underset{(-5,190)***}{0,253}\, pdoth_t - \underset{(-3,178)***}{0,036}\, pwa_t - \underset{(-16,388)***}{3,265} + D_{2,t}$$

$$ec_{3,t} = pdequ_t - \underset{(-9,279)***}{0,466}\, pdoth_t - \underset{(-1,518)}{0,018}\, pwa_t - \underset{(-12,235)***}{2,508} + D_{3,t}$$

Note: t values between brackets. *** p<0.01, ** p<0.05, * p<0.1.
Source: Own calculations.

In the second sub-period (2003-2011) the optimal lag length was found to be 10; while the cointegration test were carried out by considering an intercept and a deterministic trend (T), and including also the nominal exchange rate (*E*). The null hypothesis that the cointegration order is 4 was not rejected. The estimated cointegrating relations are given by:

$$ec_{1,t} = pdfb_t - \underset{(-12,799)***}{0,279}\, pwa_t - \underset{(-24,258)***}{1,754}\, E_t - \underset{(-14,875)***}{1,850} - \underset{(-7,587)***}{0,004}\, T$$

$$ec_{2,t} = pdclo_t - \underset{(-7,352)***}{0,102}\, pwa_t - \underset{(-26,629)***}{1,208}\, E_t - \underset{(-37,220)***}{2,944} - \underset{(-21,387)***}{0,008}\, T$$

$$ec_{3,t} = pdequ_t - \underset{(-2,372)**}{0,038}\, pwa_t - \underset{(-29,027)***}{1,539}\, E_t - \underset{(-35,850)***}{3,270} - \underset{(-7,001)***}{0,003}\, T$$

$$ec_{4,t} = pdoth_t - \underset{(-3,353)***}{0,048}\, pwa_t - \underset{(-26,855)***}{1,271}\, E_t - \underset{(-43,574)***}{3,548} - \underset{(-11,777)***}{0,004}\, T$$

Note: t values between brackets. *** p<0.01, ** p<0.05, * p<0.1.
Source: Own calculations.

The long-run elasticities coefficients were measured for the whole period 1992-2011 in an attempt to compute a global representative measure. In this case the optimal lag length is 3. The cointegration test was run by considering an intercept and seasonal dummies, an including also the exchange rate in the model. The null given by H_0: $r \leq 3$ was not be rejected. The resulting estimated equations are:

$$ec_{1,t} = pdfb_t - \underset{(-19,489)***}{1,118}\ pdoth_t - \underset{(-4,841)***}{0,236}\ pwa_t - \underset{(-3,728)***}{0,128}\ E_t + \underset{(8,440)***}{1,888} + D_{1,t}$$

$$ec_{2,t} = pdclo_t - \underset{(-6,599)***}{0,794}\ pdoth_t - \underset{(-3,691)***}{0,377}\ pwa_t - \underset{(-4,545)***}{0,328}\ E_t + \underset{(2,794)***}{1,311} + D_{2,t}$$

$$ec_{3,t} = pdequ_t - \underset{(-10,604)***}{0,777}\ pdoth_t - \underset{(-3,690)***}{0,230}\ pwa_t - \underset{(-6,349)***}{0,279}\ E_t + \underset{(0,536)}{0,153} + D_{3,t}$$

Note: t values between brackets. *** p<0.01, ** p<0.05, * p<0.1.
Source: Own calculations.

Given that the deterministic time trend was significantly different from zero when the estimation is carried out for the second sub period, an additional VEC model was adjusted by including a deterministic time trend for 1992-2011. In this case, the optimal lag length (in levels) was equal to 3, whereas the cointegrating test was carried out by considering an intercept, seasonal dummies and the time trend. The null hypothesis that suggests at least 3 cointegrating equations was not be rejected. The estimated cointegrating equations that describe the long-run relationships between domestic prices and the international price of agricultural commodities are:

$$ec_{1,t} = pdfb_t - \underset{(-15,847)***}{1,333}\ pdoth_t - \underset{(-1,709)*}{0,077}\ pwa_t - \underset{(-2,353)**}{0,073}\ E_t + \underset{(8,049)***}{2,102} + \underset{(2,870)***}{0,001}\ T + D_{1,t}$$

$$ec_{2,t} = pdclo_t - \underset{(-8,895)***}{1,202}\ pdoth_t - \underset{(-1,287)}{0,093}\ pwa_t - \underset{(-4,574)***}{0,227}\ E_t + \underset{(4,259)***}{1,786} + \underset{(3,359)***}{0,003}\ T + D_{2,t}$$

$$ec_{3,t} = pdequ_t - \underset{(-33,934)***}{1,058}\ pdoth_t - \underset{(-1,514)}{0,025}\ pwa_t - \underset{(-18,095)***}{0,208}\ E_t + \underset{(4,602)***}{0,445} + \underset{(10,103)***}{0,002}\ T + D_{3,t}$$

Note: t values between brackets. *** p<0.01, ** p<0.05, * p<0.1.
Source: Own calculations.

By comparing the two estimations for the whole period 1992-2011, it can be seen that the results remain quite sensitive to the existence of a time trend.

All in all, the different estimates suggest that the pattern of transmission of international prices into domestic prices depends upon the period under review. The differences are explained not only by changes in the applied macroeconomic policies but also by the variations in the data generating process. In the simulations below, and since our interest is to analyze the effects of the recent increase in world commodity prices, we will use the elasticities obtained for the sub-period 2003-2011. Another reason, in this case more technical, is that the sub-period 2003-2011 is the only one for which we cannot reject the null of four cointegrating relationships, so that the order in which domestic prices enter into the vector p_t is irrelevant for the estimation of their elasticities with respect to *paw*.

Equation (2) was estimated using a pool of cross-section household surveys. Table 2 reports the results. As we can appreciate from the reading of the Table there is a positive and significant relationship between the world price of agricultural commodities and the wage rate of the least skilled workers (incomplete high school or less). The same is true for those with an intermediate education, but in this case the elasticities are less than a half of those obtained for the previous group. Finally, for people with a complete tertiary or university education the estimates are in all cases not significant.

We also test if the wage elasticities are statistically different among the three groups. In all cases we can reject the null that the elasticities for the least skilled are the same as for the other two groups. When we compare those with an intermediate and upper levels of education, in three out of five cases we also reject the null that the coefficients are the same. In the simulations carried out in the next section we use the estimates from model 4.

V.2 Simulation of Welfare effects

Using the elasticities obtained before, budget shares from the National Survey of Household Expenditures (ENGHo) of 2004/2005, and assuming a 100% increase in the international price of agricultural commodities, applying equations (3) and (3') we can calculate the effect on welfare for each household under different behavioral assumptions about consumption adjustments. Then, we run non-parametric regressions of the welfare effects as a function of household per capita expenditure.[10] Figures 8 to 11 show these simulations.

Given that for all consumption categories we obtain positive elasticities of domestic prices with respect to agricultural commodity prices, it is not a surprise that all households lose with an increase in the prices of agricultural commodities. From Figure 8 we have that on the consumption side households at the lowest end of the expenditure distribution are the ones that are most affected through the increase in prices of food and beverages, while for non-food and beverages goods the opposite result arises, but the magnitude of the effects is much smaller than for food and beverages. Allowing for second order effects, as expected reduce the welfare loss, in about a 10%. There is almost no difference when we consider only own-price elasticities or both own- and cross-price elasticities. The improvement when allowing for consumption adjustments comes almost completely from food and beverages. In the aggregate (see Figure 9), are the poorer households the ones most affected by the increase in agricultural commodity prices, with losses of up to almost 12% of the household initial expenditure when no consumption adjustment is assumed, and up to 10% when we allow for households to change their consumption basket.

To obtain the income labor effects, we use the wage elasticities reported in Table 2. Then, using the income share of each member of the household, 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 10, there is a positive effect working through the increase in labor income, with this effect benefiting the most to middle income

[10] The sample sizes used in the regressions are 26431 observations. Each observation is weighted by the inverse of the probability of the household to be included into the sample.

households, and the least to the richest households. This result could be explained due to the pattern of factor intensity of Argentina's production, especially those of food exports which are intensive in the use of agricultural commodities, and most likely also intensive in semi-skilled labor. However, the increase in labor income is not enough to compensate for the welfare loss that works through the consumption of goods.[11]

Once we add the effects that work through consumption and labor income, poorest households are the most affected (see Figure 11). However, all households lose with the increase of agricultural commodity prices, the losses range from as much as 6.5% to around a 10% of the initial expenditure if no consumption adjustment is allowed, and between 6% and 9% when we allow for households to change their consumption patterns. The distribution of losses along the per capita expenditure of households is, a priori, in line with what could a priori 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.

V.4 Simulating the effect on poverty

To grasp an approximate idea of how important is the impact on poverty of an increase in world commodity prices, in Table 3 we report the indigence and poverty rates that would follow after a 100% increase in world prices of agricultural commodities, as well as two additional measures: the indigence/poverty gap and severity.[12] To get the new indigence rate, we calculate the new indigence line under the assumption that the extra amount of expenditure a household needs in order to avoid to be classified as indigent is given by the effect that works through the increase in the domestic prices of food and beverages. In the case of the poverty rate, the extra amount of money a household needs to avoid to be poor is given by the effects of changes in the domestic prices of four categories of goods: food and beverages, clothing, housing and transport and communications.[13] New household incomes are calculated taking into account only the effect of labor income of salaried household members. From the results reported in Table 3 we have that in absolute values there is an increase in indigence of 2.4 pp., while in the case of poverty it is of 3.5 pp.,

[11] The fact that we are not considering non-labor income may bias our results against the richer households, since the increase in the world prices of agricultural commodities meant an important improvement in the rent of land used in agricultural production. Also, the use of this rent by land-owners meant an important contributions to others sectors of the economy, especially the building sector.

[12] The rate, gap and severity of indigence and poverty are measured following Foster *et al*. (1984), using the following formula: $R = \frac{1}{N}\sum_{h=1}^{N}\left(\frac{z_h - y_h}{z_h}\right)^{\alpha} I^*\left(y_h < z_h\right)$, 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 $I^*(y_h < z_h)$ 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.

[13] Ideally, it would be more appropriate to work with the changes in the consumption prices of the goods that constitute the baskets of indigence and poverty. However we do not have access to price indices with that level of detail.

however while for the case of the poverty the relative increase is about 13.4%, for indigence the new value is almost 34% higher than the original figure. These changes means that about 250 thousand new households would have fallen into indigence, while 365 thousand into poverty. When we allow for households to adjust their consumption patterns, the increases in indigence and poverty rates are, respectively, 0.3 pp. and 0.5 pp. less than when there is no adjustment.

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 poverty increase more than their corresponding rate, whilst in the case of indigence the opposite outcome arises. This result means that in the case of poverty, there is not only an increase in response to the raise in the price of agricultural commodities, but also that those who were already poor as well as those who become poor, move in average further away from the poverty threshold. These results mean also that poor and indigent households become a less heterogeneous group.

V.4 A simulated policy response: elimination of the VAT in food and beverages

A policy measure that has been long asked for by part of the political and social forces, is the elimination of the Value Added Tax (VAT) on the consumption of food and beverages. In Argentina, the general VAT rate is at 21%, which means an incidence around a 17% of the price paid by the consumers. In Figure 12 we show the results of two alternative scenarios, an scenario (a) in which we assume that when the VAT is eliminated on all consumption of food and beverages, consumer prices fall 12%, with the difference captured by the producers and sellers in the form of higher profits, and an scenario (b) in which all the incidence goes to reduce the consumer prices. To keep things simple, we do not consider any price effect on other goods than food and beverages. The results of these two scenarios show that because of the larger participation of food and beverages in total expenditure, poorer households benefit more than the richer ones.

Comparing the results of Figures 11 and 12, the elimination of VAT on food and beverages would compensate for an important share of the negative effect produced by the increase in the world prices of agricultural commodities. For low and medium-low income households, under scenario (b) and assuming no consumption adjustment, the elimination of the VAT compensates almost completely for the losses. If we allow for changes in consumption patterns, the elimination of the VAT on F&B brings a welfare gain larger than the negative impact produced by the rise in commodity prices. For high-income households, the elimination of the VAT covers almost all the losses. Needless to say, this option, besides to be a one-time policy, would mean an important loss of resources for the Federal Government, which under no circumstance is an issue to be taken slightly at the moment of evaluate its feasibility, economic and politically.

VI. Summary and conclusions

The increase in the price of agricultural commodities benefited greatly to Argentina, especially in a period when the country was almost completely excluded (forcibly and/or

voluntarily) from international financial markets. On the other hand, with a large share of the population with low and medium-low incomes, the increase in agricultural commodities prices has the potential to hurt an important part of the population through a raise in the price of goods that explain an important share of households expenditures, especially those that constitute the food-basket. The evidence shows that this can be expected to be the case. A less obvious channel works through changes in factor incomes. In the case of labor income, this effect would be more beneficial to the middle income households.

Table 1
Decomposition of Argentina's export growth

		1992-2002	2002-2012
All sectors	Value	110%	217%
	Price	-9%	100%
	Quantity	130%	58%
Agricultural primary products	Value	51%	271%
	Price	-9%	139%
	Quantity	66%	55%
Manufactures of agricultural origin	Value	69%	238%
	Price	-19%	154%
	Quantity	107%	33%
Manufactures of industrial origin	Value	169%	264%
	Price	-12%	48%
	Quantity	207%	146%
Fuel and energy	Value	329%	41%
	Price	18%	350%
	Quantity	262%	-69%
Terms of trade		8%	42%

Source: own based on National Institute of Statistics and Census

Table 2
Wage elasticities

	(1)	(2)	(3)	(4)	(5)
Incomplete High School or less (a)	0.0911***	0.0859***	0.0551***	0.0945***	0.0823***
	(0.017)	(0.014)	(0.018)	(0.016)	(0.015)
Complete H. School / Incomplete Tertiary/University (b)	0.0347**	0.0332**	0.0099	0.0353***	0.0337**
	(0.017)	(0.014)	(0.014)	(0.013)	(0.017)
Complete Tertiary/University (c)	-0.0239	-0.0157	-0.0086	0.0229	-0.0141
	(0.025)	(0.022)	(0.023)	(0.023)	(0.022)
Observations	782,039	782,039	782,039	782,039	782,039
R-squared	0.844	0.846	0.847	0.847	0.847
Test of equality of coefficients (P. values)					
H_o: (a) = (b)	0.0000	0.0000	0.0000	0.0000	0.0000
H_o: (a) = (c)	0.0000	0.0000	0.0024	0.0010	0.0000
H_o: (b) = (c)	0.0000	0.0004	0.3652	0.5344	0.0002

Other explanatory variables:

Model 1: age and age squared; dummy variables for males, head of household, not-single status, education (incomplete tertiary/university and complete tertiary/university education - the reference group is incomplete secondary or less), formal job, firm size (6 to 50, and more than 50 employees - the reference group is 1 to 5 employees), type of firm (private sector, and other sector - the reference group is public sector), sector of activity (20 dummies - the reference sector is agriculture, hunting, forestry and fishing), place of residence (31 dummies - the reference city is Buenos Aires), year (14 dummies - the reference year is 2012), quarter (3 dummies - the reference quarter is January-March)

Model 2: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with a linear trend.

Model 3: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with a linear and quadratic trends.

Model 4: same as Model 1 plus the interactions of the dummy variables (except for the year and quarter dummies) with two linear trends, the first starting at the beginning of the period and the second starting in 2004.

Model 5: same as Model 1 plus the interaction of all variables (except for the year and quarter dummies) with a dummy variable equal to 1 from 2004 to 2012.

All models include a constant term and the nominal exchange rate interacted with the three educational dummies.

Robust standard errors between brackets

*** p<0.01, ** p<0.05, * p<0.1.

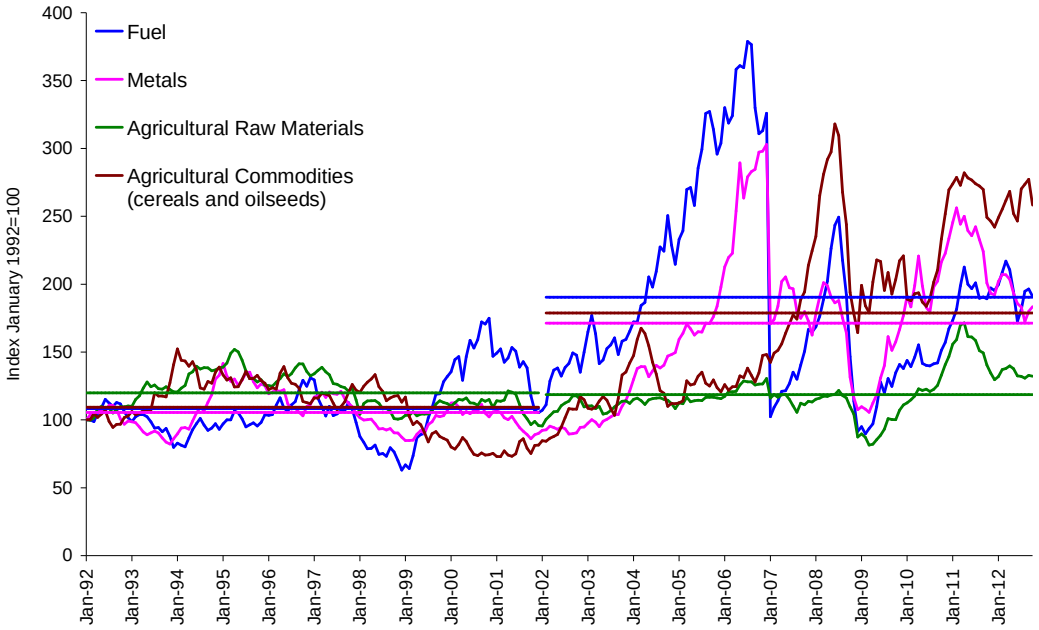
Table 3
Indigence and Poverty rates in urban areas
Pre and post a 100% increase in world prices of agricultural commodities

	Variable	Adjustment in consumption		
		No	Yes (1)	Yes (2)
Indigence (A)	Rate (pre)	7.1%	7.1%	7.1%
	Rate (post)	9.5%	9.1%	9.1%
	Gap (pre)	2.4%	2.4%	2.4%
	Gap (post)	3.2%	3.0%	3.0%
	Severity (pre)	1.3%	1.3%	1.3%
	Severity (post)	1.6%	1.5%	1.5%
Poverty (B)	Rate (pre)	26.0%	26.0%	26.0%
	Rate (post)	29.4%	28.9%	28.9%
	Gap (pre)	9.9%	9.9%	9.9%
	Gap (post)	11.6%	11.3%	11.3%
	Severity (pre)	5.3%	5.3%	5.3%
	Severity (post)	6.3%	6.1%	6.1%

(A) For the post values only includes the effects of changes in the prices of food and beverages. (B) For the post values only includes the effects of changes in the prices of food and beverages, clothing, housing, and transport and communications. (1) Considering only own-price elasticities. (2) Considering own- and cross-price elasticities.

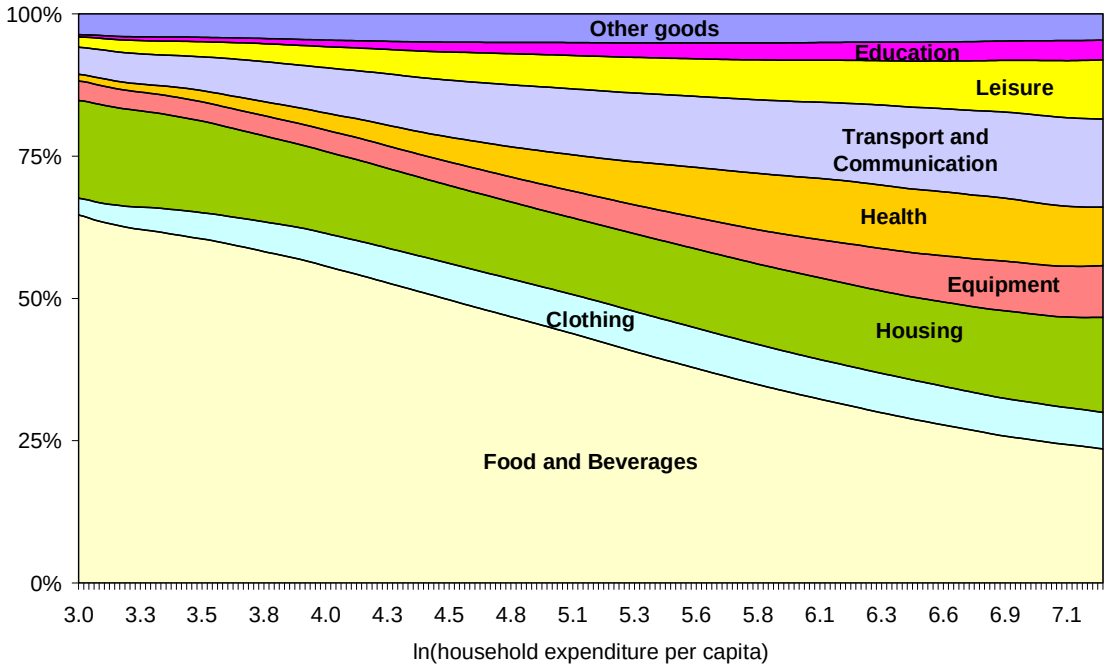
Note: simulated (post) values for columns 2 and 3 differ between them at the second or third decimal.

Figure 1
Evolution of main primary commodity prices

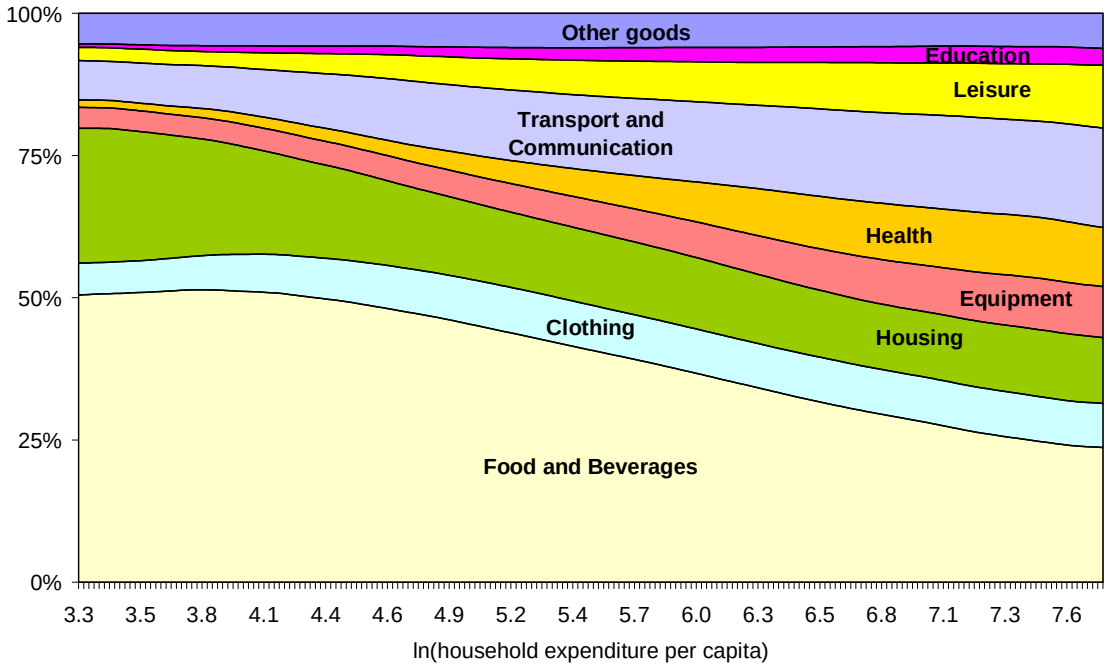


Source: own based on WITS and www.indexmundi.com (retrieved on November 12, 2012).
Note: for agricultural commodities is an export weighted average of the world prices of Argentina's main commodities

Figure 2
Expenditure shares and household expenditure (*)
1996/1997

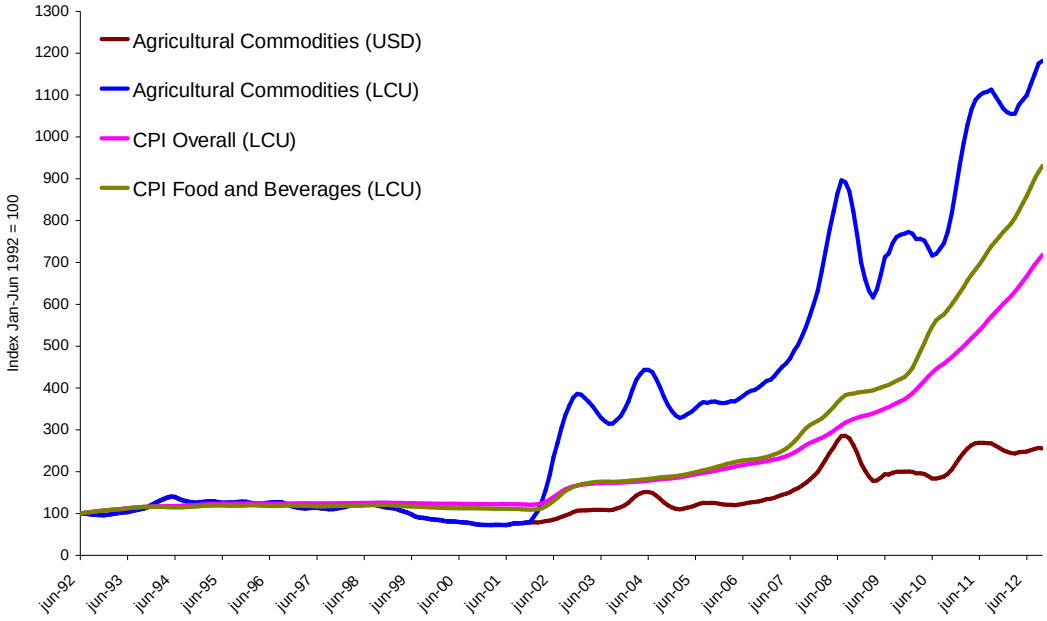


2004/2005



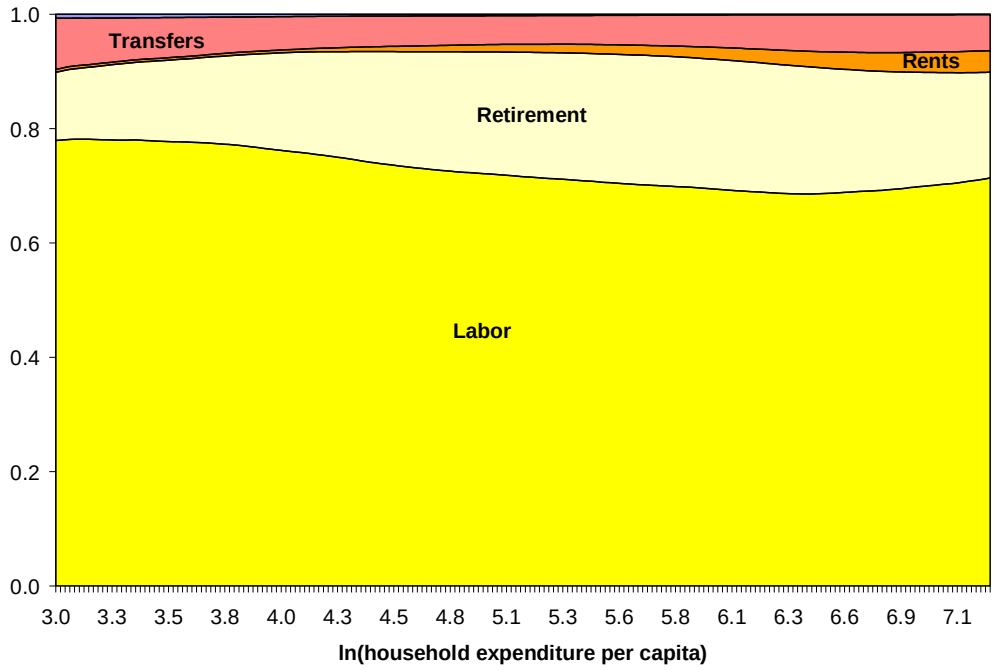
(*) The relationships between expenditure shares and expenditure per capita were obtained by non-parametric regressions.
Source: own based on ENGHo 1996/1997 and ENGHo 2004/2005.

Figure 3
Agricultural commodity and consumer prices
(six-month moving average)

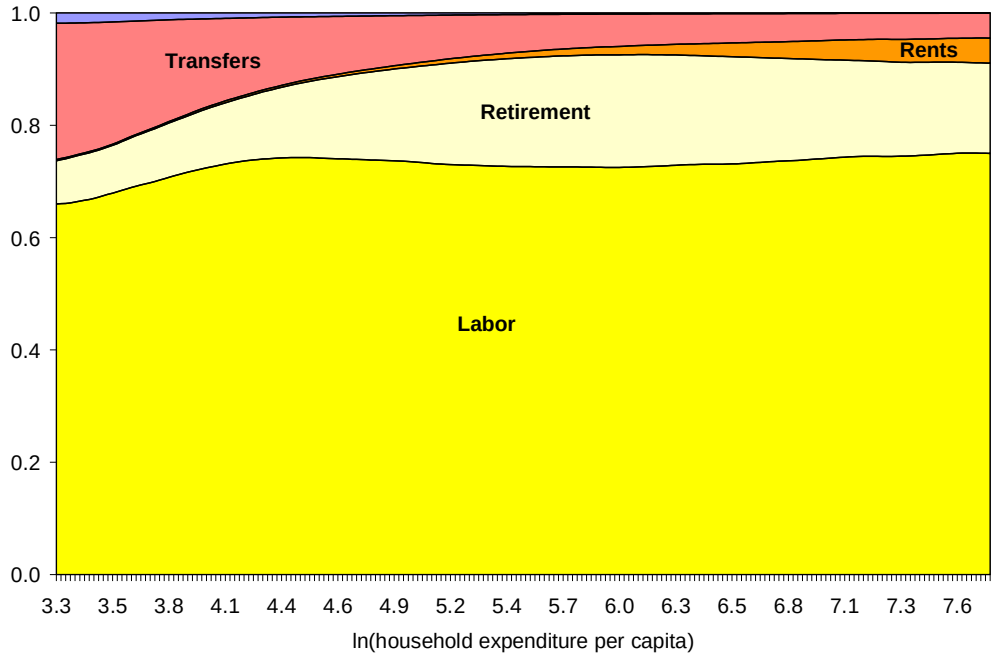


Source: own based on WITS, Instituto Nacional de Estadísticas y Censos, Government Province of San Luis, and www.indexmundi.com (retrieved on November 12, 2012)

Figure 4
Income sources and household expenditure (*)
1996/1997

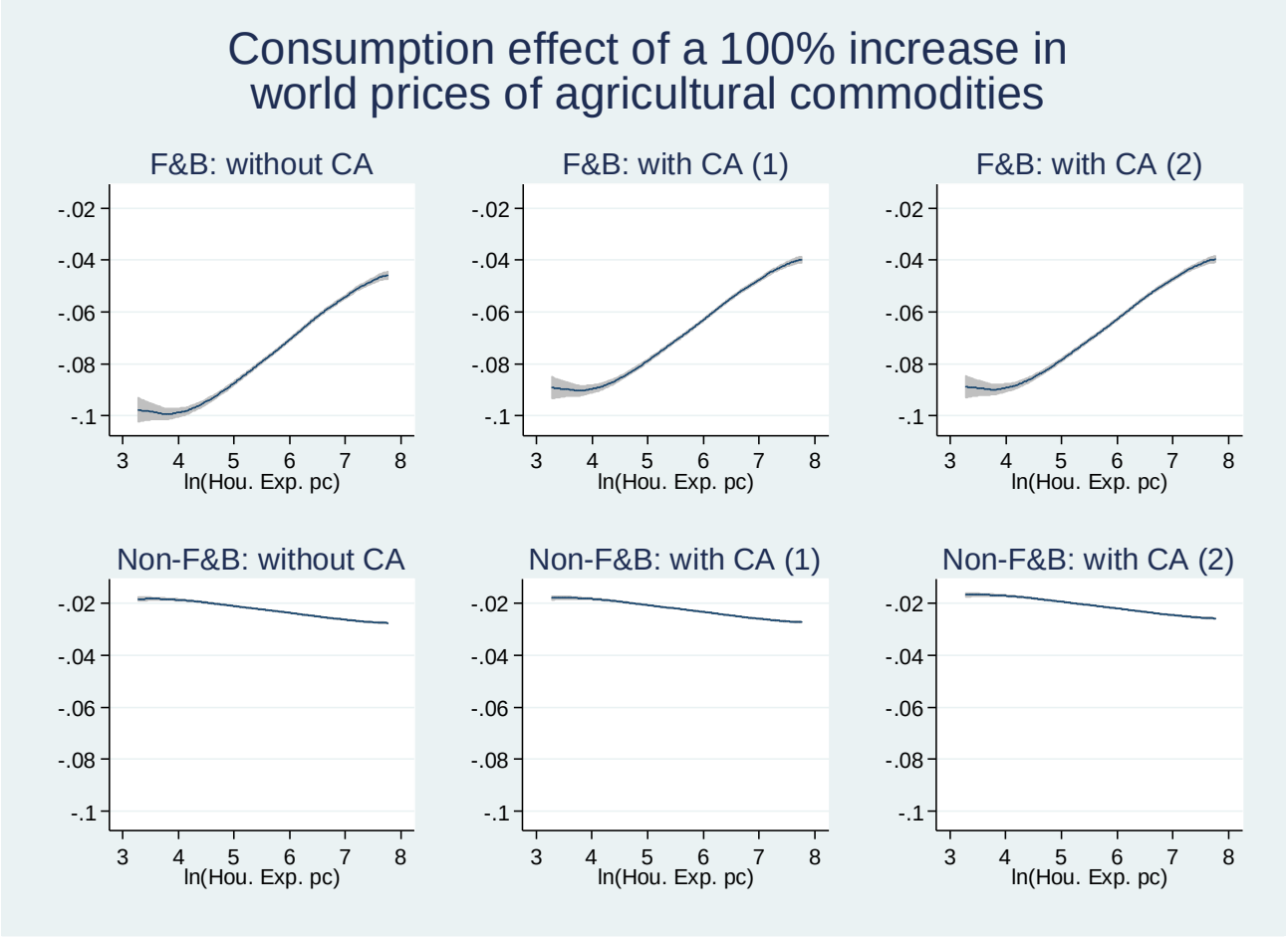


2004/2005



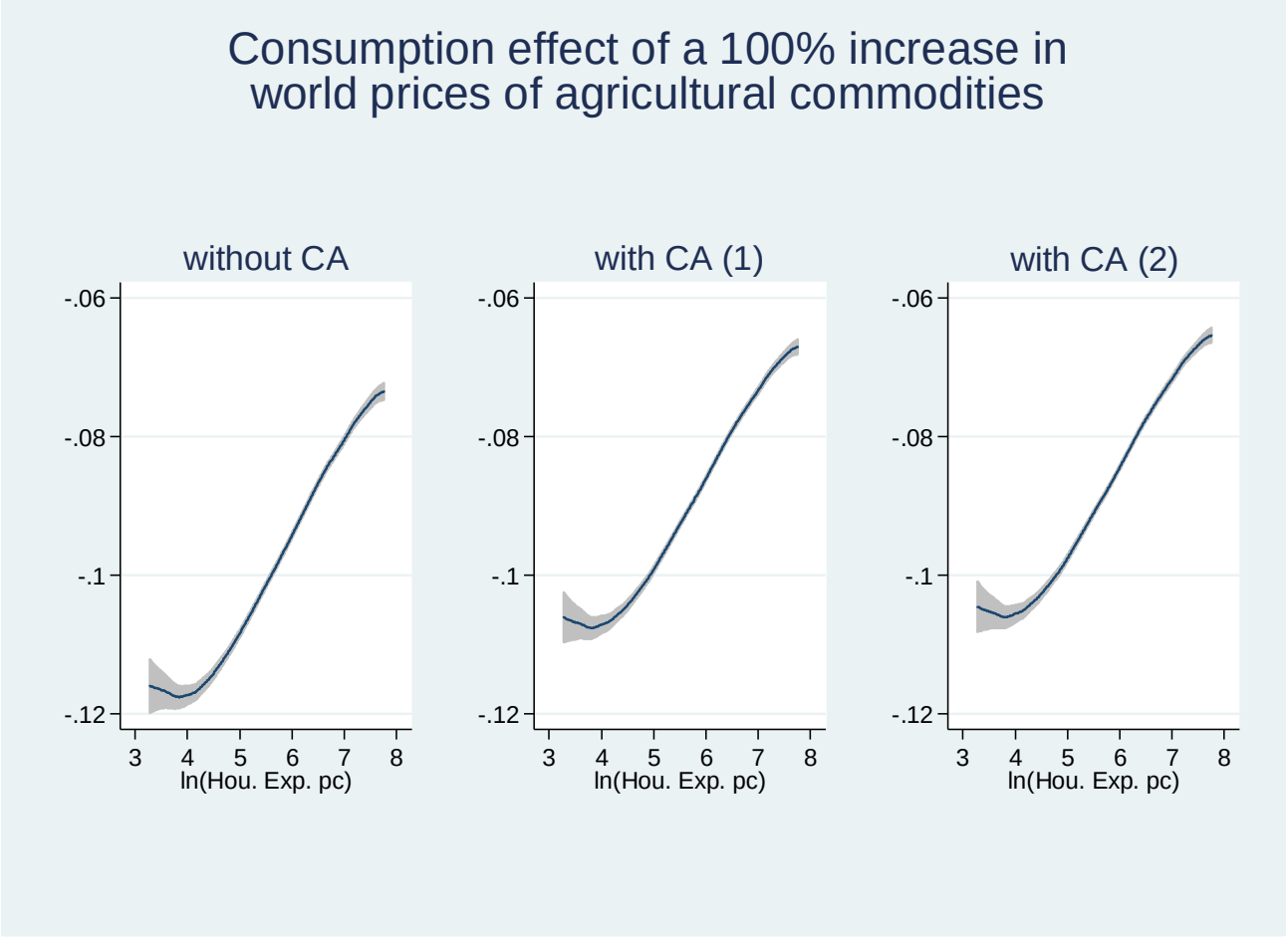
(*) The relationships between income shares and expenditure per capita were obtained by non-parametric regressions.
Note: Labor includes salaried workers, self-employed and employers; Rents include housing rents, dividends and interest.
Source: own based on ENGHo 1996/1997 and ENGHo 2004/2005.

Figure 8



(1) considering only own-price elasticities. (2) considering own- and cross-price elasticities.

Figure 9



(1) considering only own-price elasticities. (2) considering own- and cross-price elasticities.

Figure 10

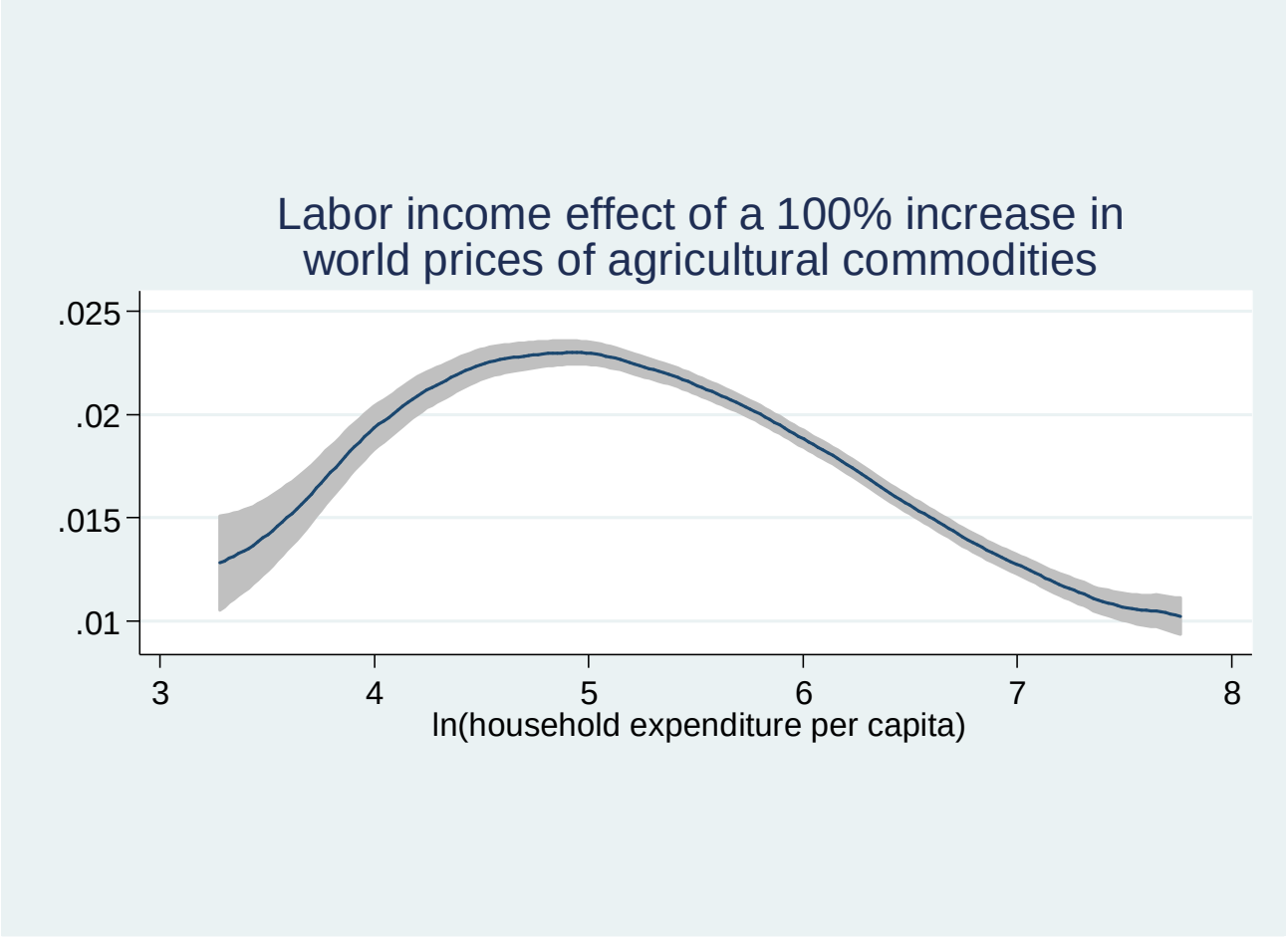
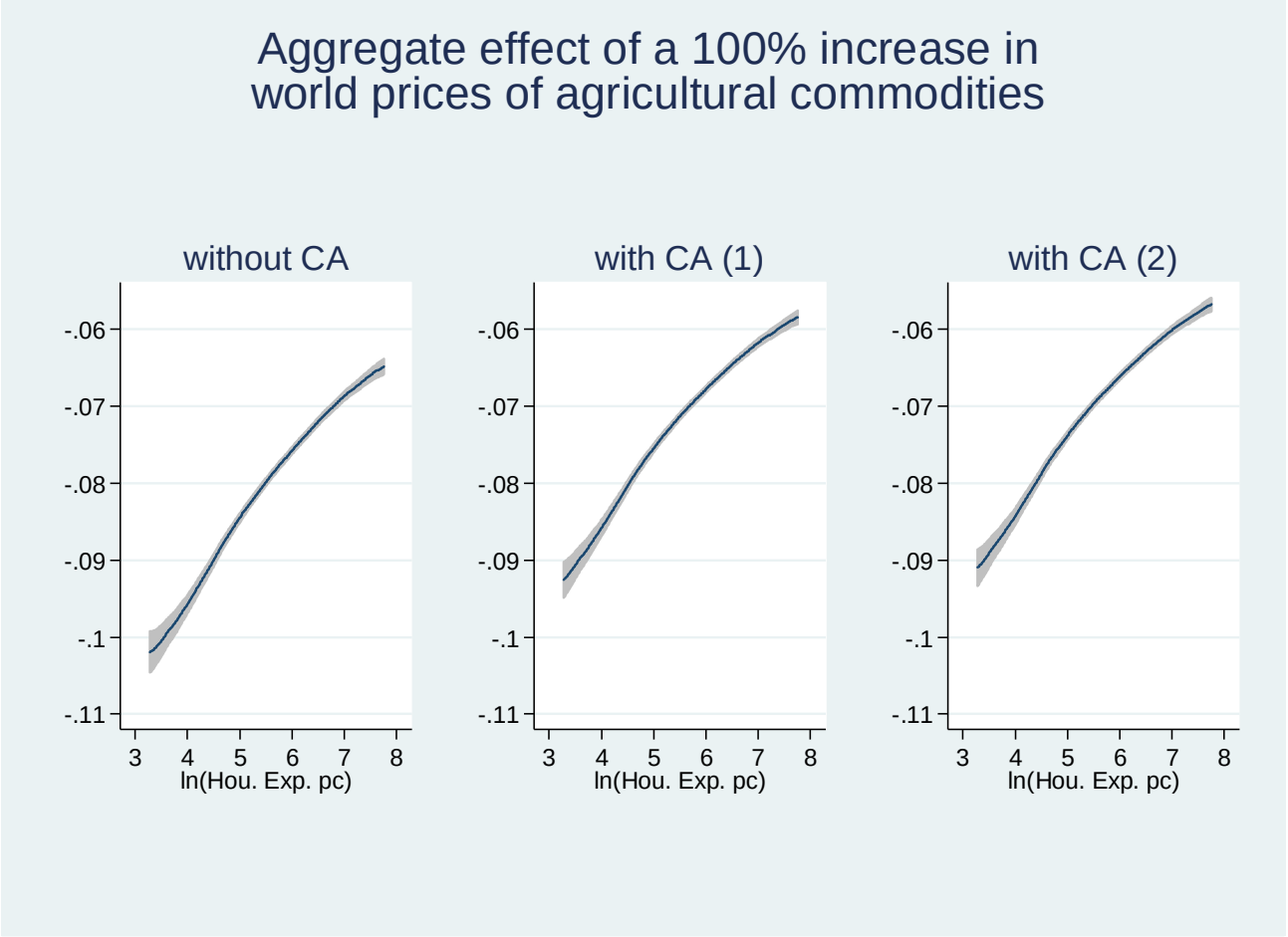
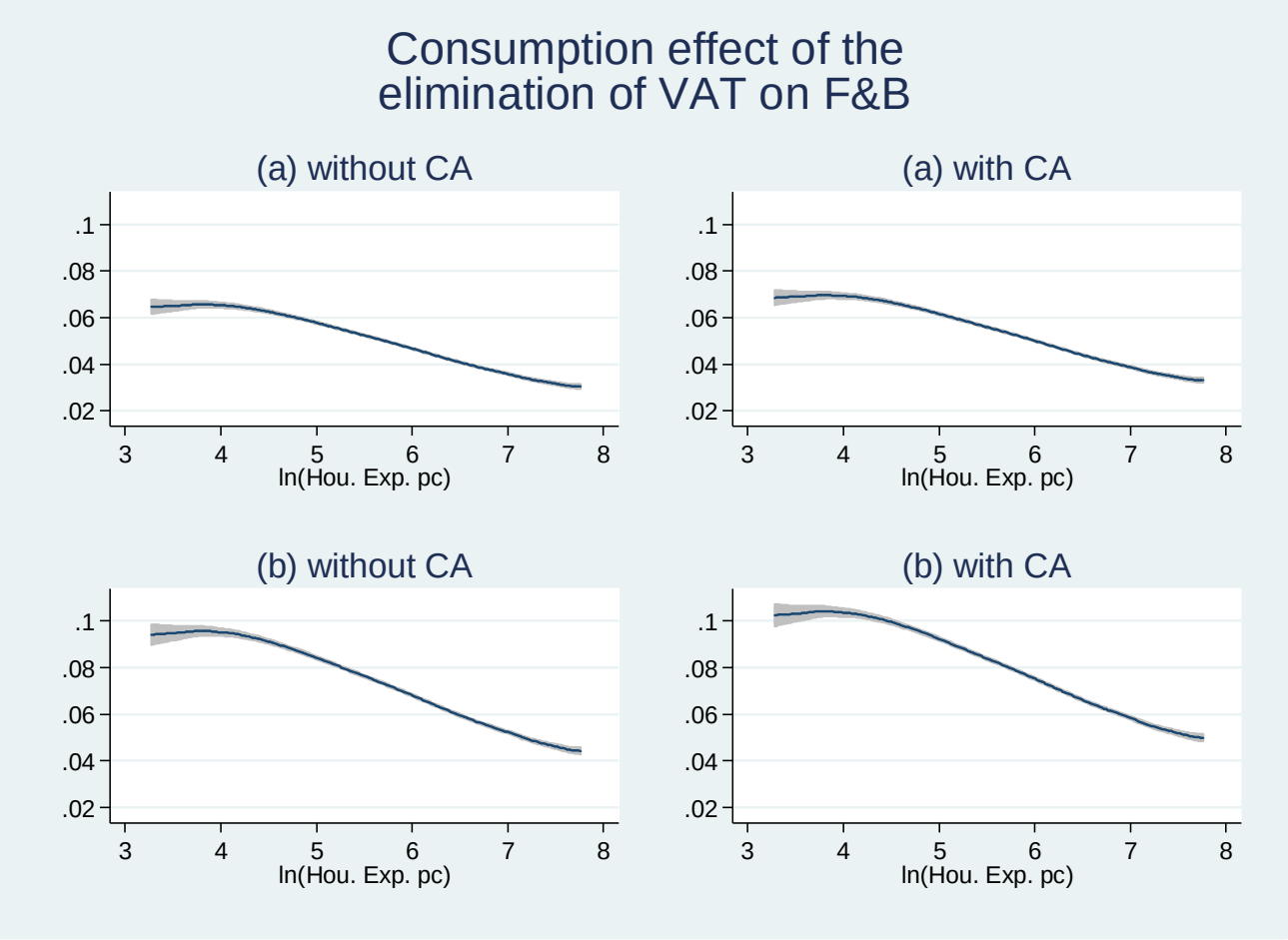

Labor income effect of a 100% increase in world prices of agricultural commodities
.025
.02
.015
.01
3
4
5
6
7
8
ln(household expenditure per capita)

Figure 11



(1) considering only own-price elasticities. (2) considering own- and cross-price elasticities.

Figure 12



(a) Prices of F&B falls by 12%. (b) Prices of F&B falls by 17%.

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Appendix:

Data sources

Nominal Exchange Rate	Banco Central de la República Argentina
Exports	WITS (World Integrated Trade Solution) of World Bank
Agricultural Commodity Index: weighted average of the prices of Maize, Soybeans, Wheat, Soybean Oil, and Sunflower Oil. Argentina's exports are used as weights	Own based on www.indexmundi.com and WITS
Soybeans: U.S. soybeans, Chicago Soybean futures contract (first contract forward) No. 2 yellow and par, US Dollars per Metric Ton	www.indexmundi.com (retrieved on November 12, 2012)
Soybean Meal: Chicago Soybean Meal Futures (first contract forward) Minimum 48 percent protein, US Dollars per Metric Ton	
Soybean Oil: Chicago Soybean Oil Futures (first contract forward) exchange approved grades, US Dollars per Metric Ton	
Maize (corn): U.S. No.2 Yellow, FOB Gulf of Mexico, U.S. price, US Dollars per Metric Ton	
Sunflower Oil: US export price from Gulf of Mexico, US Dollars per Metric Ton	
Wheat, No.1 Hard Red Winter, ordinary protein, FOB Gulf of Mexico, US Dollars per Metric Ton	
Commodity Fuel index:includes Crude oil (petroleum), Natural Gas, and Coal Price Indices	
Metals Price Index: includes Copper, Aluminum, Iron Ore, Tin, Nickel, Zinc, Lead, and Uranium Price Indices	
Agricultural Raw Materials Index: includes Timber, Cotton, Wool, Rubber, and Hides Price Indices	
Consumer Price Indices	Instituto Nacional de Estadísticas y Censos and Government Province of San Luis
Household Expenditure Survey (Encuesta Nacional de Gastos de los Hogares) 1996/1997 and 2004/2005	Instituto Nacional de Estadísticas y Censos
Household Survey (Encuesta Permanente de Hogares) 1995 to 2011	