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Poverty and Agricultural Growth: Chile in the 1990s

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Abstract

This paper analyzes the roles of agriculture in reducing poverty. Following the methodology proposed by López (2002), three channels by which agricultural growth reduces poverty are tested: (i) its effects on the real wage of unskilled workers (and/or its possible effect in reducing their unemployment); (ii) the direct impact of agricultural growth on the income of poor farmers; and, (iii) the effect on real food prices. The paper concludes that the pro-poor role of agricultural expansion is dramatic. Agricultural growth tends to improve all measures of poverty significantly with head count falling around 7.3% as a consequence of a 4.5% increase in agricultural output. An important result is that while the economy-wide effects taking place via food prices and especially the labour market are quantitatively important the direct income effects on farmers are almost negligible.

***Keywords**: agricultural growth, Chile, poverty, rural development*

1. Introduction

The objective of this paper is to quantitatively analyze the roles of agriculture in reducing poverty. Following the methodology proposed by López (2002), we consider three channels by which agricultural growth can affect poverty, namely its effects on the real wage of unskilled workers (and/or its possible effect in reducing their unemployment), the direct impact of agricultural growth upon the income of poor farmers and its effect on real food prices. The latter effect is mostly relevant for food

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items that are not traded. That is, for commodities that due to large transaction costs or to government policies have endogenous domestic prices.

We first econometrically estimate the impact of agricultural growth on the real wage of unskilled workers, its employment level, real food prices and income of poor farmers. Next we simulate how such changes impact the income of the poor and how they affect the extent and depth of poverty. We consider a benchmark case based on the current situation and then using the same methodology we simulate how poverty would have changed if agricultural and agro-processing output had been higher than it actually was. Also, we consider how increasing the share of agriculture and agro-processing in total GDP would affect poverty. The idea is to test the hypothesis that a shift in the composition of national output toward agriculture (keeping total national output constant) is pro-poor and to measure the quantitative importance of such effect if the hypothesis is not rejected.

The remainder of the paper is organized as follows: Section 1 provides a general overview of the evolution of poverty in Chile over the last decade and a half. This section also discusses general economic trends. Section 2 presents the econometric estimates of the labour demand equations which are the basis for most of the exercise. Section 3 reports on the estimates of food prices and some estimates of the contribution of agriculture to the income of poor farmers. In each section we present poverty simulations separating each of the three channels mentioned above. We also present the consolidated effects of all mechanisms on extent and depth of poverty. Section 4 concludes.

2. Poverty Profile (1987-2000)

During the 1990s the Chilean economy underwent a period of fast economic expansion which translated into a successful poverty reduction experience. As shown in Table 1, except for the last two years of the century that showed a slow-down, per capita GDP experienced rapid growth. Agricultural growth, however, slowed down from 1994 onward.

Table 1 Income growth: national and agricultural related sectors (annual %)

	1987/90	1990/92	1992/94	1994/96	1996/98	1998/2000
Per Capita GDP Growth	5.4	8.3	4.6	7.4	4.2	0.8
Agricultural Per Capita GDP Growth	7.3	4.5	2.7	1.7	-0.1	0.7
Agriculture + Food Processing Per Capita GDP Growth	5.7	6.8	3.3	3.2	-0.4	N/A

Source: Banco Central de Chile

As seen in Table 2, most poverty measures show a sharp reduction. We focus in the FGT class of poverty indicators:

$$FGT(\alpha) = \sum_{i \in P} [(z - y_i)/z]^{\alpha} / N ,$$

where z is the poverty line, y_i is the household per capita income of each person, N is total population, and the summation is over all the persons that are poor, *i.e.* with per capita income below the poverty line[1]. FGT(0) refers to the incidence of poverty, or headcount ratio, the ratio of those that are poor to total population. FGT(1), also known as the poverty gap, measures both the incidence of poverty as

[1] In this section we focus in monetary measures of poverty. World Bank (2002) studies non monetary measures of poverty and also found important improvements in these types of indicators.

well as its depth. The income gap measures as a percentage the gap of the average income of the poor with respect to the poverty line. It is straightforward to show that this income gap is equal to FGT(1)/FGT(0). Finally, the FGT(2) measure, is a more "poverty averse" indicator since it weights more heavily those persons with a higher gap as it sums the square of the income gap of those that are poor.

Table 2 Poverty and inequality

Total Population	1987	1990	1992	1994	1996	1998	2000
Poverty							
Head Count – FGT(0)	46.08	38.46	32.43	27.4	23.13	21.62	20.58
Income Gap	42.55	38.36	34.93	35.31	33.65	34.48	34.52
Poverty Gap – FGT(1)	19.61	14.75	11.33	9.67	7.78	7.46	7.1
FGT(2)	11.18	7.86	5.57	4.96	3.83	3.78	3.71
Inequality							
GINI	56.74	0.5523	54.97	55.18	55.22	55.88	55.83
Coefficient of Variability	1.6789	1.8636	1.8027	3.3318	1.7495	1.8556	1.9319
Rural Population							
% Rural	19.47	18.55	18.04	16.54	16.12	14.57	14.14
Poverty							
Head Count – FGT(0)	53.47	39.37	33.37	30.8	30.57	27.57	23.82
Income Gap	40.15	37.06	32.42	33.5	33.37	32.88	34.62
Poverty Gap – FGT(1)	21.47	14.59	10.82	10.32	10.2	9.06	8.25
FGT(2)	11.52	7.83	5.03	5.11	4.95	4.33	4.28
Inequality							
GINI	49.02	57.73	50.69	50.38	49.27	49.69	51.08
Coefficient of Variability	2.0213	2.6651	2.1475	2.2339	1.7090	3.1124	2.8564

Source: Authors' calculations using CASEN surveys.

We measure poverty using the official poverty line developed by ECLAC and the Chilean Ministry of Planning (Table 2). Additionally, for international comparison purposes in Table 3 we measure poverty using a standard poverty line equivalent to 2 US dollars per day. In the following discussion we centre in the first two indicators. The official poverty line is equal to 2 minimum food baskets for urban areas. A separate food basket is calculated for rural areas, where food prices are lower, and the poverty line is calculated as 1.75 times the value of the rural basket of food, as it is assumed that the cost of services is lower in rural areas.

Table 3 Internationally comparable poverty measures. 2 US dollars per day poverty line

Total Population	1987	1990	1992	1994	1996	1998	2000
Head Count – FGT(0)	43.62	33.22	26.04	23.91	20.23	18.79	17.57
Income Gap	42.70	37.53	33.74	34.49	33.96	34.70	34.95
Poverty Gap – FGT(1)	18.62	12.46	8.79	8.25	6.87	6.52	6.14
FGT(2)	10.61	6.55	4.25	4.23	3.41	3.33	3.25

Note: A 2 US dollars of 1994 per day line was used for the above calculations. Source: Authors' calculations using CASEN surveys

At the beginning of the period less than 1/5 of the population was considered rural, and by the end of the period the proportion of rural population fell to about 1/6. This is in part due to a change of definition of rural locations that took place in 1996. Prior to 1996, locations with less than 2000 inhabitants were considered rural. From 1996 onwards the definition of rural changed to locations with less than 1,000 inhabitants or between 1,000 and 2,000 where less than 50% of the economically active population is employed in secondary (industry) and tertiary (services) activities. Thus, rural numbers are not strictly comparable between 1987-1994 and 1996-2000[2].

As shown in Table 2, the incidence of poverty shows a marked reduction; in less than a decade from 1987 to 1996 it was halved from 46% to 23%. The poverty gap and the FGT(2) indicators have also shown similarly impressive reductions during the period. However, looking at the income gap, one can see that this indicator fell during the 1987-1992 period, after that it remained relatively stable at around 34-35%. This indicates that the reductions of the FGT(1) and FGT(2) indicators after 1992 are due mainly to reductions of the incidence of poverty, and not to reductions in its depth. Rural poverty, on the other hand has shown a similarly fast reduction pattern. Except for the period 1994-1996, rural poverty has fallen throughout the 1990s even during the last two years of the century when the economy overall slowed down.

Although it has been a stated objective of the governments of the period, there have not been important improvements in inequality. The more broadly used GINI coefficient shows a slight improvement from 1987 to 1992, but followed by an increase in inequality from 1994 onwards. The coefficient of variability that applies a greater weight to distances in the upper tail, that is to the richest, shows a steady increase in the period. The exception is 1994, with an extremely low inequality level that may probably be due to a mis-sampling of the wealthiest in the 1994 CASEN survey.

3. Estimating the wage, price and direct income effects

The Structure of Labour Demand

We postulate that producers in the economy minimize the cost of production. There are two outputs being produced, agriculture and agro-processing (Q_a), and everything else (Q_n). These outputs are produced using three variable factors of production, unskilled labour (L_u), skilled labour (L_s), and capital (K). Producers are assumed to be competitive facing exogenously given factor prices (w_u for unskilled labour, w_s for skilled labour and w_r for the rental price of capital). The three factors of production are also assumed to be mobile across the two productive sectors and are allocated in a way that their marginal products are equalized across the sectors. Under these assumptions cost minimization implies that there exists an aggregate dual cost function, $C(w_u, w_s, w_r; Q_a, Q_n; t)$, where t

[2] Also, note that Chile uses a rather narrow definition of rural compared, for example, to the 10,000 inhabitants threshold used in Switzerland.

stands for the level of technology. The cost function must satisfy certain properties: It is non-decreasing, linearly homogeneous and concave in factor prices, and non-decreasing in each of the outputs (Diewert 1971).

We specify a multioutput Generalized Leontief (GL) form for the cost function, which is considered a flexible functional form (Diewert 1971). GL is flexible because it does not impose *a priori* any form of separability neither any restrictions on the elasticities of substitution among the factors of production. This is, in contrast for example with the highly popular Cobb-Douglas function, which *a priori* imposes that the marginal rate of substitution of all input pairs are independent of other inputs (separability) and that all elasticities of substitution are equal to one. Similarly, the CES function imposes all the above restrictions except that it allows for non-unitary but identical elasticities substitution among all input pairs. Imposing these restrictions reduces the reliability and usefulness of empirical estimates that rely on functional forms that are not flexible.

A specification for the multioutput GL is the following:

$$C = Q_a \sum_i \sum_j b_{ij} (w_i w_j)^{1/2} + Q_n \sum_i \sum_j c_{ij} (w_i w_j)^{1/2} + tQ_a \sum_i b_i w_i + tQ_n \sum_i c_i w_i + Q_a Q_n \sum_i d_i w_i \,, \tag{1}$$

where b_{ij}, c_{ij}, b_i, c_i, and d_i (i,j = u,s,r) are coefficients. The GL cost function *a priori* only imposes linear homogeneity in factor prices. All other properties of the cost function could in principle be tested. The implicit demand equations for unskilled and skilled labour are derived from (1) using Shephard's lemma,

$$L_s = \sum_j b_{sj} (w_j / w_s)^{1/2} Q_a + \sum_j c_{sj} (w_j / w_s)^{1/2} Q_n + b_s t Q_a + c_s t Q_n + d_s Q_a Q_n \,, \tag{2}$$

$$L_u = \sum_j b_{uj} (w_j / w_u)^{1/2} Q_a + \sum_j c_{uj} (w_j / w_u)^{1/2} Q_n + b_u t Q_a + c_u t Q_n + d_u Q_a Q_n \,. \tag{3}$$

The linear homogeneity property of $C(\bullet)$ in (1) leads to factor demands that are homogenous of degree zero in all factor prices. Also using Shephard's lemma one could derive the demand for capital. However, we only have data for the price of capital but not for its quantity demanded. Therefore, we omit the equation for capital demand although we do use its rental price on the right hand side of labour demand equations. Equations (2) and (3) can be jointly estimated after imposing the symmetry conditions, $b_{ij} = b_{ji}$ and $c_{ij} = c_{ji}$.

Estimation Procedure

The labour demand equations (2) and (3) are estimated using data from CASEN surveys for the period 1990-96. We have data for each one of the 13 regions of the country and for each one of the four years in which the survey was implemented (1990, 1992, 1994, 1996). That is, we have a regional data panel of 52 observations (13 regions × 4 periods) for each of the equation variables, except for the rental price of capital for which we have only annual prices for the whole country (but see below for an explanation of why this price can still be used in the panel estimates). We have data of total unskilled labour and skilled labour used in each region over the period, also from the CASEN survey we estimated wages for each one of the two types of labour. The distinction between skilled and unskilled labour was based on years of schooling; workers with schooling of less than 8 years (primary and compulsory school in Chile) were considered "unskilled" while workers with greater schooling are labelled "skilled." From other sources we were able to calculate regional GDP separated between agriculture and non-agricultural industries.

Also from non CASEN sources we obtained implicit unit values of imported capital goods. The unit values for imported capital goods were used as proxies for the domestic price of capital.[3] These prices do not vary across regions and, in fact, the price of capital is not likely to change much across regions of the country. Despite the lack of variability across regions we still have enough variability of

[3] Please see the appendix for more details regarding the data.

the *relative* factor prices to allow us to estimate cross demand elasticities, e.g., elasticities of labour demand with respect to the price of capital. This is due to the homogeneity condition derived from the theoretical restrictions arising from the assumption of cost minimization. As can be seen in (2) and (3), the two labour demand equations are explained by factor price ratios. The rental price of capital appears as w_r/w_s in (2) and w_r/w_u in (3). These ratios vary not only over time but, given that w_s and w_u change across regions as well, also across regions.

Table 4 shows the estimated coefficients for equations (2) and (3). In estimating these equations we have used a Two-Stage Iterated Seemingly Unrelated Regression procedure which is equivalent to Maximum Likelihood estimates. As can be seen in the table, the goodness-of-fit of the system is highly satisfactory as shown by the high level of significance of most coefficients. Moreover the estimated coefficients are consistent with several of the properties of a cost function. The estimated cost function is monotonically increasing and concave in factor prices, and increasing in each one of the outputs.

Table 4 Estimates of the multioutput cost function. Restricted, Iterated 2S-SUR Method

	Coefficient	Standard Error		Coefficient	Standard Error
b11	299.136	372.316	b21	0.637	1.969
b12	0.637	1.969	b22	1497.927**	335.746
b13	-13.840	88.674	b23	-527.454**	111.630
c11	-28.881	45.179	c21	0.016	0.317
c12	0.016	0.317	c22	-280.571**	40.215
c13	10.374	10.092	c23	112.550**	12.958
b1	-0.149	0.186	b2	-0.745**	0.167
c1	0.014	0.023	c2	0.140**	0.020
d1	-6.73E-08	3.11E-07	d2	1.73E-06**	2.50E-07

Note: One asterisk denotes significance at the 10% level, while 2 asterisks denote significance at the 5% level. Source: Authors' calculations.

An issue of potential concern is the use of a pseudo panel as done in this study. When cohorts averages (regions in our case) are used instead of true panel data there is room for inconsistency in the estimators as we use sample averages (random by nature) instead of true means. However, it has been determined that when the cohort sample is large enough (larger than 100 observations), this bias tapers off (see Verbeek and Nijman, 1992).[4]

An even bigger concern arises from the use of wages as exogenous variables in our estimations when economic theory suggests that both employment and wages are endogenously determined. We address this issue by using predicted (square root) wage ratios in a two-stage procedure. We first predicted wage ratios using several combinations of the following instruments: regional population, agricultural and non-agricultural output, and average labour schooling. In the second stage we used the predicted regional wage ratios in the joint estimation of equations 2 and 3.

Our biggest concern for system misspecification arises from measurement errors in the wage variables used. Implicitly in the calculation of the wage rates we are assuming that a day worked in metropolitan Santiago is equal to a day worked in mostly rural region XI. If the wage differentials observed are in part determined by different labour efforts, there is the risk of bias in the price

[4] In our case we feel comfortable that the variance of our sample means is small enough to eliminate any serious inconsistency as the smallest regional sample is larger than 2,400 observations.

elasticities of our estimates. Although it is hard to define the size of this bias, the functional specification chosen, which relies on wage ratios, mitigates this bias as long as the direction of the bias is the same for each labour type in each different region (a likely scenario). Also, the fact that we use instrumental variables may reduce the effect of measurement errors.

The Elasticities: Effects of Changes in Agricultural Output Level
Table 5 shows the elasticities of demand for unskilled and skilled labour implicit in the estimated coefficients and evaluated at sample means. It also presents the standard errors of these elasticities (note that the elasticities are functions of several coefficients) and their degree of statistical significance. As can be seen the two labour demand equations are downward sloping with similar own price elasticities of the order of -0.53 to -0.61[5]. All own price demand elasticities are statistically significant at least at 5%. Unskilled and skilled labour are substitute inputs as shown by the fact that the cross price elasticities are both positive; the elasticity of unskilled labour demand with respect to skilled labour wage rate is about 0.21 while the elasticity of skilled labour demand with respect to the unskilled wage rate is about 0.07. However, these cross elasticities are not individually statistically significant but they are jointly significant.

Table 5 Estimated labour demand elasticities (evaluated at sample means)

	Unskilled Labour	Skilled Labour	Capital	Agricultural Output	Non Agricultural Output
Unskilled Labour	-0.53** (0.2669)	0.21 (0.3911)	0.32 (0.4242)	0.58*** (0.1071)	0.40*** (0.0728)
Skilled Labour	0.07 (0.681)	-0.61*** (0.2045)	0.54** (0.2454)	0.44*** (0.0581)	0.70*** (0.0384)

Note: Standard errors in parentheses. *** Significant at the 1% level. ** Significant at the 5% level. *Significant at the 10% level. Source: Authors' calculations. *** Significant at the 1% level. ** Significant at the 5% level. *Significant at the 10% level. Source: Authors' calculations.

Unskilled labour appears to be substitute with capital, with an elasticity of labour demand with respect to the rental price of capital of 0.32. Skilled workers on the other hand, are clearly substitutes with capital with an elasticity of skilled labour demand with respect to the rental price of capital of about 0.54. This suggests that the concern so often found in the literature about the potential displacement of labour by capital does not appear to be empirically grounded for the case of unskilled labour, but rather surprisingly may be apparently valid for skilled labour.

The most important empirical finding is the asymmetric response of the two types of labour to output expansion in the agricultural and non-agricultural production. The demand for unskilled workers is much more sensitive to an expansion of agriculture than skilled workers. In fact, the elasticity of unskilled labour demand with respect to agricultural output is almost 0.6 compared to an elasticity for skilled labour of only 0.44. Additionally, unskilled labour demand is more responsive to agricultural output than to non-agricultural output; while skilled labour exhibits opposite responses. The demand for unskilled labour is less than 60% as responsive to non-agricultural output as the demand for skilled labour is. As shown in Table 5, all four labour demand elasticities with respect to

[5] This value is highly consistent with previous estimates of (own price) labour demand elasticities obtained by studies for Chile that have usually ranged between -0.4 and -1 (see for example, Fajnzylber and Maloney (2000) and Riveros (1985) .

output are statistically significant at 1% level. These results imply that increasing the share of agriculture in total output while keeping total output constant would lead to an expansion of employment of unskilled workers. That is, agricultural based economic growth is more favourable for unskilled (usually poor) workers than economic growth with a stagnant agricultural sector.

Table 5 also suggests that the output scale elasticities (e.g., the sum of the elasticities with respect to Q_a and Q_n) are remarkably close to one for the case of unskilled labour and about 1.1 for the case of skilled workers. In fact the output scale elasticities are not statistically different from one. This suggests a very close link between overall economic growth and labour demand with skilled labour demand being only slightly more responsive than unskilled labour demand. Thus, the apparent difficulties of the economy to create jobs are not necessarily related to low labour/output intensities.

Effect of Changes in Output Mix or Composition

To estimate the effect of a change in the *composition* of production towards the agricultural sector we proceed as follows: Total output can be approximated by the following aggregator,

$$Q = Q_a^{\gamma} Q_n^{1-\gamma},$$

where γ is the weight of agriculture and agro-processing in aggregate output and 1 - γ is, of course, the corresponding share of all other sectors. Logarithmic differentiation yields,

$$d\ln Q = \gamma d\ln Q_a + (1-\gamma) d\ln Q_n.$$

To consider the pure composition effect we allow a compensated change of Q_a so that Q does not change. This means that $d\ln Q = 0$. Hence, we get that to maintain total output constant the expansion of agriculture has to be compensated by a corresponding fall of Q_n. Hence,

$$\left.\frac{d\ln Q_n}{d\ln Q_a}\right|_{dQ=0} = \frac{-\gamma}{1-\gamma}.$$

Thus, the compensated logarithmic effect of Q_a on L_u is,

$$\left.\frac{d\ln L_u}{d\ln Q_a}\right|_{dQ=0} = \frac{\partial \ln L_u}{\partial \ln Q_a} + \frac{\partial \ln L_u}{\partial \ln Q_n} \left.\frac{\partial \ln Q_n}{\partial \ln Q_a}\right|_{dQ=0} \tag{4}$$

From our estimates (Table 5) we have that $\frac{\partial \ln L_u}{\partial \ln Q_a} = 0.58$ and $\frac{\partial \ln L_u}{\partial \ln Q_n} = 0.40$. Also the share of agriculture/agro-processing in total GDP (γ) over the period was about 0.15. Hence,

$$\left.\frac{d\ln L_u}{d\ln Q_a}\right|_{dQ=0} = 0.58 - 0.40 \cdot \frac{0.15}{0.85} = 0.51.$$

That is, a 1% compensated increase in agricultural output (which needs a 0.15/0.85 = 0.17% fall in non-agricultural output) induces an increase in the demand for unskilled workers of the order of 0.5%. In other words, a 1% increase in the share of agriculture keeping total output constant raises the demand for unskilled workers by 0.5%, while raising skilled labour demand by only 0.3%

Simulating Poverty Impacts of Agriculture through Labour Demand Mechanisms: The Assumptions

In this section we use the previous results to evaluate the impact of agricultural growth on poverty as measured by the headcount index and the income gap of the poor relative to the poverty line. We do not have reliable estimates of unskilled labour supply elasticities for Chile. We have been able to find only one study for Chile that estimates a labour supply elasticity (Mizala, Romaguera and Henríquez 1999). This means that we have little information to compare and evaluate the reliability of the findings in this study. For this reason we alternatively use two extreme assumptions that give the boundaries of these effects, and next we contrast them using the Mizala et al. estimate for the labour supply elasticity. In the first extreme case we assume that labour supply elasticity is zero, that is, supply of labour is fixed in which case only the wage rate changes. At the other extreme labour supply is assumed completely elastic, in which case the wage rate is fixed and, as a consequence, the labour market may not clear and unemployment may prevail. In this second case the full adjustment is absorbed by employment change. This latter assumption is consistent, for example, with labour market regulations or binding minimum wages. Finally, we use the Mizala estimates of labour supply elasticity (1.8), which allows for wages and employment to adjust simultaneously.

(a) Fixed supply of unskilled workers. If the supply of unskilled labour is fixed, and further assuming that the rental price of capital also remains fixed, (an assumption consistent with the open economy assumption) then:

$$dL_i = \frac{\partial L_i}{\partial w_u} \cdot dw_u + \frac{\partial L_i}{\partial w_s} \cdot dw_s + \frac{\partial L_i}{\partial Q_a} \cdot dQ_a = 0,\ i = u, s \qquad (5).$$

Dividing (5) by L_i and rearranging terms, we can express the two equations as a system with two unknowns:

$$\begin{bmatrix} \varepsilon_{u,u} & \varepsilon_{u,s} \\ \varepsilon_{s,u} & \varepsilon_{s,s} \end{bmatrix} \cdot \begin{matrix} d\ln w_u / d\ln Q_a \\ d\ln w_s / d\ln Q_a \end{matrix} = -\begin{bmatrix} \varepsilon_{u,Q_a} \\ \varepsilon_{s,Q_a} \end{bmatrix} \qquad (6),$$

where $\varepsilon_{i,i}$ refers to labour demand elasticities, for example $\varepsilon_{u,s} \equiv \frac{\partial L_u}{\partial w_s} \cdot \frac{w_s}{L_u}$. Using the labour demand elasticities presented in Table 5, we find that the unskilled wage elasticity with respect to agricultural output is $d\ln w_u / d\ln Q_a = 1.43$, as the wage rate adjusts to clear the labour market. That is, a 1% increase in agricultural output, *ceteris paribus*, will cause an increase in the wage rate of the unskilled of 1.43%.

The elasticity of the skilled wage with respect to agricultural output is estimated in a similar way. Given the estimated elasticities in Table 5, we have that $d\ln w_s / d\ln Q_a = 0.89.$ That is, the effect of agricultural output on the skilled wage rate is only about 60% the value of the corresponding agricultural output elasticity on unskilled labour.

The compensated wage effect of agricultural/agro-processing output keeping total output in the economy constant can be estimated as with (6), but replacing the right hand side of the equation ($-\varepsilon_{i,Q_a}$) with $-\varepsilon_{i,Q_a} - \varepsilon_{i,Q_n} \cdot \left(d\ln Q_n / d\ln Q_a \right)\Big|_{dQ=0}$; that is, the (negative) compensated labour output elasticities as calculated above. Solving (6) with the compensated output elasticities we find that the compensated unskilled wage elasticity with respect to agricultural output is 1.21 (0.66 for skilled wages). That is, a change in the composition of GDP that increases the share of agriculture/agro-processing by 1% leads to an increase in the wage rate of the unskilled of 1.2%.

(b) Fully elastic supply of unskilled workers. If the supply elasticity of labour is elastic instead of inelastic as assumed above the wage effect of agricultural growth is smaller. In fact the larger is the labour supply elasticity, *ceteris paribus*, the smaller is the wage effect. The other bound for the poverty

effect occurs when the labour supply elasticity is infinity in which case we only have an employment effect instead of a wage/employment effect. In this case the effect is an increase in unskilled employment with an elasticity equal to 0.57, while the compensated employment effect is 0.51.

(c) The intermediate case. If the labour supply elasticity is neither 0 nor infinity, but a known constant, we can equate the change in labour demand in (5), to the change in labour supply:

$dL_i^s = \frac{\partial L_i^s(\cdot)}{\partial w_i} \cdot dw_i, i = u, s$, where L_i^s is labour supply for skill level *i*. Again, dividing by L_i and rearranging terms, we can express the two equations as a system with two unknowns:

$$\begin{bmatrix} \varepsilon_{u,u} - \eta_{u,u} & \varepsilon_{u,s} \\ \varepsilon_{s,u} & \varepsilon_{s,s} - \eta_{s,s} \end{bmatrix} \cdot \begin{bmatrix} d\ln w_u / d\ln Q_a \\ d\ln w_s / d\ln Q_a \end{bmatrix} = -\begin{bmatrix} \varepsilon_{u,Q_a} \\ \varepsilon_{s,Q_a} \end{bmatrix} \qquad (7),$$

where $\eta_{i,i}$ is the own price elasticity of labour supply, assumed to be identical for skilled and unskilled labour, and equal to 1.8, as estimated by Mizala, Romaguera and Henríquez, 1999. The solution of (7) indicates that the wage elasticities with respect to agricultural output are 0.26 and 0.19 for unskilled and skilled labour respectively. In addition, in this case there is also an employment effect,

$$\Delta \ln L_i = \frac{\partial \ln L_i^s}{\partial \ln w_i} \frac{d\ln w_i}{d\ln Q_a} \Delta \ln Q_a .$$

Simulation Results

From the previous analysis we estimated the elasticities of unskilled wage rates and employment with respect to agricultural growth. It was assumed that all unskilled and skilled workers benefited of their corresponding wage increase, which is translated into higher per capita household income. The increased employment effect on the other hand, was simulated by adding the average wage rate per skill level to the incomes of those that are economically unemployed, that are able to become employed. The amount of individuals that become "employed" in such manner is determined by the corresponding labour elasticities. Also, each additional individual "employed" will increase household income which may either lead the household out of poverty (e.g., the number of poor declines by the number of individuals in the household) or the increased household income may not be sufficient to bring its per capita income above poverty or the household may be already above the poverty line. Each new worker was randomly selected from the pool of unemployed. Thus, the household per capita income increases through these two channels, namely wage and employment effect. With this new household income profile we performed the new estimates for head count poverty, poverty gap, etc.

The first column of Table 6 provides the benchmark poverty situation in Chile in the year 2000. It includes headcount poverty, the income gap of the poor, and also includes the proportion of the vulnerable population, defined as those whose income is less than the poverty line plus 20%, as well as their mean proportional income surplus above that line. The next three vertical blocks in Table 6 show the simulation results under the three alternative assumptions for the case where agricultural output increases 4.5%, which has been the actual growth rate until the late 1990s. Within each block we present the uncompensated and compensated (constant aggregate output) effects. Blocks I and III show the simulation results under alternative extreme assumptions of fixed supply of unskilled labour (in which case the effect of agricultural growth is reflected on wage increases only) and an infinitely elastic supply consistent with unemployment caused, for example, by a binding minimum wage (in which case the effect of agricultural growth occurs via greater employment). In Block II we present the intermediate case that assumes a labour supply elasticity of 1.8 obtained from Mizala, Romaguera and

Henríques (1999) (in which case agricultural expansion causes both higher wages and greater levels of employment).

Surprisingly we discover that the reduction of poverty incidence brought about by agricultural expansion is equivalent for the both the extreme assumptions about labour, infinitely elastic and inelastic labour supply. This is a fortuitous result and should not, however, be considered a rule. Furthermore, although the reduction in the head-count ratio is equivalent, Table 5 shows that the case of employment effects only (column III) is much more egalitarian, with stronger reductions in the "poverty averse" indicators FGT(1) and FGT(2), as well as in the income gap. As indicated in Table 6, an uncompensated expansion of agricultural output reduces head-count poverty from 20.6% to about 19.2%, regardless of the assumption about labour supply elasticity, or equivalently a reduction of 6.6%. The income gap is reduced but only very slightly from 0.1% in the fixed labour supply case to about 2.3% in the other two cases. Poverty vulnerability is also reduced slightly in all three cases while the income surplus of the non-poor but vulnerable groups also marginally increases.

The compensated effects (e.g., the case when there is only a change in the composition of output in form of agriculture but not an expansion of aggregate output) follow similar patterns, but of course are of a smaller magnitude. Head-count also falls within a very narrow range around 5.2%. In any case these effects are still quite large suggesting that a more agricultural based growth is clearly pro poor.

Table 6 Effects of agricultural expansion on poverty: labour market effects

	Bench-mark 2000	I Wage Effects Only†		II Wage and Employment Effects§		III Employment Effects Only‡	
		(1)	*(2)*	*(1)*	*(2)*	*(1)*	*(2)*
Poverty							
Head Count (%)	20.58	19.23	19.50	19.20	19.56	19.23	19.51
Income Gap (%)	34.52	34.48	34.44	33.64	33.82	33.75	33.86
FGT(1)	7.10	6.63	6.72	6.46	6.62	6.49	6.61
FGT(2)	3.71	3.48	3.52	3.32	3.41	3.33	3.40
Vulnerable Groups							
Poverty line + 20%							
Head Count (%)	7.34	7.25	7.24	7.11	7.16	7.18	7.23
Income Surplus (%)	9.81	10.09	10.03	9.95	9.87	9.78	9.82
Per Capita Income (100s Ch.$2000)	1,383	1,429	1,418	1,410	1,403	1,404	1,399
Income Change (%)		3.33	2.56	1.94	1.46	1.57	1.18

(1) Uncompensated Simulations. (2) Compensated Simulations. † Wage effect under the assumption that labour market clears and supply of unskilled labour is fixed, i.e. labour supply elasticity is zero. § Employment expansion and wage effect using estimated labour supply elasticity of 1.8 from Mizala et al., 1999. ‡ Employment expansion effect under the assumption that labour supply is fully elastic and unemployment prevails, i.e. labour supply elasticity is infinity. Source: Authors' calculations.

4. Food Prices and Agricultural Growth

In this section we examine the hypothesis that agricultural growth helps reducing the real price of food products that are not tradable. Evidence would suggest that in Chile most of the agricultural growth has been outward oriented with expansion of tradable goods and processed agricultural products. However, whether or not this growth has spillovers to non-tradable food prices remains an empirical question that we address analyzing time series of prices.

To determine the marginal effect of agricultural growth on food prices we explain the path of the real non-tradable food price index (*PINTF*) as a function of external factors, real exchange rate (*RER*), and internal factors: real non-food price index (*PINF*), agricultural output (Q_a), and non agricultural output (Q_n):

$$PINTF_t = \alpha + \delta \cdot t + \beta_1 RER_t + \beta_2 PINF_t + \beta_3 \ln Q_{at} + \beta_4 \ln Q_{nt} + \mu_t \qquad (8)$$

where α is a constant, and μ_t a random error term. The first problem we have to deal before estimating (4) is that some or all series are expected to be non-stationary. For example, if both agricultural output and non agricultural output have a long run growth rate, with yearly deviations from those level then by definition $\ln Q_a$ and $\ln Q_n$ are non-stationary series, integrated of order 1– I(1): $\Delta \ln Q_n = a + \eta_t$, where a is a constant and η_t is a mean zero random perturbation. The use of non-stationary series in regressions yields invalid, spurious correlation results, usually red-flagged by an R^2 approaching to 1 as sample size grows.

Thus, we run a battery of unit root tests on our series to detect the presence of integrated time series. The results presented in Table 7 suggest as expected that the log of agricultural and non-agricultural output are I(1): both the Phillips-Perron and the Advanced Dickey Fuller (ADF) tests do not reject the null hypothesis for the series in levels, but strongly reject the null for the series in first differences. There is strong evidence that the rest of the time series are also I(1). In spite of our small sample size (1976-2000), most of the tests results are strong, except for the Phillips-Perron test on the levels of the non-food real price index which is almost rejected at the 10% level, and the tests for the first differences of the real exchange rate which at the 5% level are inconclusive. The unit root tests performed on the second differences of the real exchange rate reject the null (ADF=-5.44), which would altogether suggest that *RER* is I(2). Other researchers have also found evidence of aggregate price levels being I(2) (e.g., Clements and Mizon (1991); Miller (1991)). However, it turns out that this result may be due to the low power of the ADF test. In fact, both Clements and Mizon, and Miller carry out their subsequent analysis assuming that aggregate prices are I(1). We also treat RER as I(1) given that the test results are not overwhelming, and our sample size is very small.

Table 7 Food price equation: unit root tests

Series	Levels			First Difference		
	Phillips-Perron	*ADF*	*Interpolated 5% Critical Level*	*Phillips-Perron*	*ADF*	*Interpolated 5% Critical Level*
Non-Tradable Real Food Prices	-2.813	-1.909	-3.600	-3.789	-3.962	-3.000
Real Exchange Rate	-1.330	-2.155	-3.600	-3.089	-2.977	-3.000
Non-Food Real Prices	-3.129	-2.378	-3.580	-7.454	-7.398	2.989
Log Agricultural Output	-2.857	-1.209	-3.580	-5.721	-5.676	2.989
Log Non-Agricultural Output	-1.652	-2.194	-3.580	-3.867	-3.793	2.989

Note: In both tests H_0: the series has a unit root. A trend term was used in the tests for the series in levels. 3 lags were used in both set of tests to control for autocorrelation. Source: Authors' calculations.

Assuming that all of the time series are integrated of order 1 we test for cointegration among series. We carried out full information maximum likelihood procedures, as suggested by Johansen (1991), not rejecting the hypothesis that the series cointegrate. However, here we report the Phillips and Ouliaris (1990) single equation procedure, because we attempt to only explain one series, and the power of the full information method is very low with small samples like ours. The Phillips and Ouliaris procedure amounts to test for the stationarity of the residual of (8), μ_t. The ADF and Phillips-Perron tests of -6.22 and -4.83 respectively, lie below the asymptotic critical value at 5% of -4.49.[6] Therefore, we conclude that the residual of (8) is stationary, and equivalently the time series cointegrate, with $[1 \quad \beta]$ as a cointegrating vector.

Cointegration first of all means that the results of estimating (8) are not spurious. Cointegration means that two or more series that are non-stationary, although may deviate from each other in the short run, have a long run equilibrium; *i.e.* in spite of the deviations, they move together. A classic and intuitive example of cointegration is given by aggregate consumption and income series. Both are usually non-stationary, but cointegrate as they have a long-run equilibrium, given by a long run share of income devoted to consumption. The fact that our series cointegrate indicates they too have a long-run equilibrium. That is, the real price of non-tradable food may deviate from real price of non food goods, the real exchange rate, and agricultural and non-agricultural output, but in the long-run the five move together.

The results of estimating (4) are presented in Table 8. We find some intuitive results: the real price of food is positively correlated with the real price of non food goods, the real exchange rate and the non-agricultural output. The relationship of our interest, between agricultural output and the price of non-tradable food goods is as hypothesized negative, however this relationship is not as strong as the effect of non-agricultural output with a coefficient almost significant at the 5% level. The estimated coefficient of -0.56, evaluated at the sample mean translates to a long-run elasticity of non-tradable food prices to agricultural output of -0.18.

Table 8 Non-tradable real food prices: estimated long-run effects

	Coefficient	Std. Error	t-Stat
Constant	247.16***	34.83	7.10
Trend	-0.1426***	0.0211	-6.77
Real Exchange Rate	0.0232***	0.0040	5.85
Non Food Real Prices	2.8425*	1.6155	1.76
Log Agricultural Output	-0.5620*	0.2749	-2.04
Log Non-Agricultural Output	2.7893***	0.5332	5.23

R^2=0.67. Observations: 24; 1977-2000. Std. Error of Regression = 0.4947. *** Significant at the 1% level. ** Significant at the 5% level. *Significant at the 10% level. Durbin-Watson Stat.: 1.83; Critical Values at 5% (0.925-1.902). Newey-West Standard Errors Presented.
Source: Authors' calculations.

Furthermore, as in the previous section, we can calculate the compensated elasticity of non-tradable food prices to agricultural output,

$$\left.\frac{d \ln PINTF}{d \ln Q_a}\right|_{dQ=0} = \frac{\partial \ln PINTF}{\partial \ln Q_a} + \frac{\partial \ln PINTF}{\partial \ln Q_n} \frac{\partial \ln Q_a}{\partial \ln Q_n}$$

[6] Critical value from Phillips and Ouliaris (1990), furthermore the 2.5% critical value is -4.77, and the 1% critical value is -5.04.

$$= -0.18 + 0.9 \cdot (-0.15 / 0.85) = -0.34 .$$

This compensated elasticity is higher than the uncompensated one because non-agricultural output, which is required to fall in a constant total output agricultural expansion, is strongly correlated with non tradable food prices.

We proceed to estimate the error correction representation to analyze the short-run relationships, by estimating:

$$\Delta PINTF_t = a + \gamma_1 \Delta RER_t + \gamma_2 \Delta PINF_t + \gamma_3 \Delta \ln Q_{at} + \gamma_4 \Delta \ln Q_{nt} + \lambda \hat{\mu}_{t-1} + \varepsilon_t \qquad (9)$$

where a is a constant, $\hat{\mu}_t$ is the estimated residual from (8), and ε_t is a random perturbation. In (5), the γ_i represent the short-run relationships, as opposed to the long run relationships captured in the first regression. The estimated coefficient of λ of -0.97 represents the proportion of the short-run deviation in t - 1 that is offset by a movement in $PINTF_t$; *i.e.* almost all of the departure from the long-run equilibrium is eliminated within a year. The results of estimating (9) are presented in Table 9, where we can see that the signs of the short-run relationships are equivalent to the long-run presented in Table 8. An important result is that the short-run effect of agricultural output on non-tradable food prices is much larger than the long-run effect. This may be due to the fact that when prices go up some commodities in the non-tradable food price index start being traded in the international market. That is, the price of some of these commodities may leave the range of transaction costs. But initially non-traded commodities can become tradable only after some adjustment period has elapsed. This result suggests that increases in the growth rate of agricultural output has a strong short-run (1 year) effect in decreasing the real price of non-tradable food, but this effect is reduced after one year as the long-run equilibrium is re-established. For the poverty simulations presented below we use the estimated long-run elasticities.

Table 9 Change in non-tradable real food prices (error correction representation): estimated short-run effects

	Coefficient	Std. Error	t-Stat
Constant	-0.1475***	0.0195	-7.55
Change Real Exchange Rate	0.0226***	0.0029	7.94
Change Non Food Real Prices	3.4377***	0.9098	3.78
Change Log Agricultural Output	-0.9367**	0.3984	-2.35
Change Log Non-Agricultural Output	3.2451***	0.3708	8.75
Lagged Error	-0.9665***	0.1365	-7.08

R^2=0.68. Observations: 23; 1978-2000. Std. Error of Regression = 0.4763. *** Significant at the 1% level. ** Significant at the 5% level. *Significant at the 10% level. Durbin-Watson Stat.: 1.81. Critical Values at 5% (0.895-1.920). Newey-West Standard Errors Presented. Source: Authors' calculations.

Food Price Poverty Effects: Simulations

The reduction of non tradable food prices affects poverty twofold. First, the price fall increases real incomes; at the same time the cost of the food basket that is used to measure poverty becomes lower, i.e., the poverty line falls. The change in real income from a reduction in food prices can be calculated using the weights from the national CPI which reflects consumption patterns for society as a whole. According to the consumption survey used to construct the 1998 base year CPI, 27% of household budgets are spent on food. Thus, if the price of food falls 1%, then real incomes grow by: $1/(1-1\% \cdot 27\%) - 1 = 0.3\%$.

An additional effect of food price changes is its impact upon the poverty line. The poverty line which as explained earlier is equal to 2 food baskets is of course quite sensitive to changes in the price of food. We used the same estimates to evaluate the fall in the poverty line as a consequence of a reduction of food prices.

With the information on the changes in real income and the poverty line, we simulate the effect of an expansion of agriculture by 4.5%, both by itself, and holding total output constant. Table 10 shows the result of this exercise. In the first column we show the benchmark poverty measures for the year 2000, the second column presents poverty effects of food price reductions following an expansion of the agricultural sector. The results suggest that through the price effect a 4.5% growth of agriculture would at most reduce the incidence of poverty by about ½ %, and would leave the vulnerability (as measure by the population with per capita income up to 20% above the poverty line) of those almost poor relatively unchanged.

Table 10 Effects of agricultural expansion on poverty: Food price effect

		Food Price Effects	
		(1)	*(2)*
Poverty			
Head Count (%)	20.58	20.43	20.33
Income Gap (%)	34.52	34.49	34.42
FGT(1)	7.10	7.05	7.00
FGT(2)	3.71	3.68	3.65
Vulnerable Groups			
Poverty line + 20%			
Head Count (%)	7.34	7.29	7.23
Income Surplus (%)	9.81	9.78	9.83
Per Capita Income			
(100s Ch.$2000)	1,383	1,383	1,384
Income Change (%)		0.1	0.1

(1) Uncompensated Simulations. (2) Compensated Simulations. Source: Authors' calculations.

As can be seen in Table 10, unlike the labour market effect, the price effect is larger in the compensated case than in the non-compensated one. The reason is that in the compensated case non-agricultural output falls causing a further reduction of agricultural prices (note in Table 8 that the coefficient of non-agricultural output in agricultural prices is positive).

Direct Effects on Poor Farmers

The direct poverty effect via the income of poor farmers was also measured. We found that even under the most optimistic assumptions this effect is negligible. To determine the importance of agricultural/agro-processing output in the income of poor farmers we estimated a regression where both off farm income as well as agricultural output explain total household income of poor farmers, so as to determine the relative shares of each component (Table 11). The use of this simple approach is mostly due to the fact that we lack information on key variables affecting poor farmers' income such as land holdings as well access to other assets. Nonetheless, the estimated elasticity of 0.1 seems plausible when one considers the increasing importance of non-farm employment and incomes in rural Chile reported by Berdegué, Ramírez and Reardon (2001). These authors show that the importance of non-farm income has been growing in rural areas, adding up to 41% of overall rural income. This does not mean that the share of agricultural related income in total poor farmers' income is only 10%. In fact a significant part of their remaining income is still derived from agricultural activities outside their own farm. We note that the effects of off-farm agricultural output sources on poor farmers occur essentially via wage and off-farm agricultural employment levels. These effects are ignored here because they are already accounted for in the labour market analysis.

Table 11 Estimates for the participation of agricultural income in poor farmers' per capita income

	Coefficient	Std. Error
Log Off-Farm Per Capita Income	0.9911***	0.0596
Log of Agricultural Product	0.1219**	0.0577

Observations: 31 Std. Error of Regression = 3.6921 *** Significant at the 1% level,
** Significant at the 5% level, * Significant at the 10% level
White's Robust Standard Errors Shown.

	Coefficient	Std. Error
Log Off-Farm Per Capita Income	1.013***	0.0568
Log of Agricultural Product + Ag. Processing Output	0.0984*	0.0535

R2 =0.99. Observations: 31. Std. Error of Regression = 3.7743. *** Significant at the 1% level,
** Significant at the 5% level, Significant at the 10% level. White's Robust Standard Errors Shown.
Source: Authors' calculations.

Furthermore, the incidence of subsistence farming in Chile is very limited. About 1.4% of those working are considered to be subsistence farmers in the country. As a consequence, the exercise of simulating increases in poor farmers' income as a result of agricultural growth provided a negligible effect on poverty.

Consolidated Results

We finally consolidate the poverty alleviating effects of an agricultural expansion of 4.5% in Table 12. The most striking result is that there is not a big difference between the compensated and uncompensated effect of agricultural expansion on poverty. In the uncompensated case, the labour market effects dominate for a reduction of the incidence of poverty around 7.3%. In fact the labour market effect explains about 90% of the total poverty reduction while the food price effect explains the remaining 10%. The composition change effects of agricultural expansion are also clearly pro-poor. An increase in the share of agriculture by 4.5% reduces poverty by about 6.4%. In this latter case, however, the price effects play a more important role, explaining about 20% of total poverty reduction.

Table 12 Consolidated effects of agricultural growth of 4.5% on poverty

	Wage Effects Only	Wage and Employment Effects	Employment Effects Only
Uncompensated	-7.29	-7.43	-7.29
Compensated	-6.46	-6.17	-6.41

Source: Authors' calculations.

5. Conclusion

This paper evaluates the role of agricultural growth in reducing poverty in Chile. The analysis is broad enough to allow for economy-wide mechanisms including wage changes and food price changes that affect poverty among both rural and urban households. The paper also measures the direct impact of agricultural expansion on the income of farmers that are poor and near poor. An important result is that while the economy-wide effects taking place via food prices and especially the labour market are quantitatively important the direct income effects on farmers are almost negligible.

The pro-poor role of agricultural expansion appears to be quite dramatic: Agricultural growth tends to improve all measures of poverty significantly with head count falling around 7.3% as a consequence of a 4.5% increase in agricultural output. That is, the elasticity of poverty reduction with respect to agricultural growth falls within the range of 1.6%, substantially larger than elasticities normally found for aggregate growth in Chile which is of the order of 0.8 to 1.2 (Contreras 2002; World Bank 2002). That is, agricultural growth has not only a large impact on poverty but also its effect is much greater than the effect of expanding other sectors in the economy.

This latter result is corroborated by our finding that the *compensated* effect of agricultural growth is also positive and large. In fact, a 4.5% increase in agricultural output, keeping total output constant, leads to poverty reductions in the range of 6.4%. That is, the compensated elasticity of poverty reduction is about 1.4. Interestingly, our estimates are highly consistent with the aggregate output elasticities of around 1 found by Contreras. Given that the share of agriculture plus agro-processing in national outputs is about 0.15 we have that a 1% compensated rise in agricultural output means the rest of the economy must contract by about 0.17%. If the aggregate poverty elasticity is about 1 as estimated by other studies, the 0.17% contraction by the non-agricultural sector should cause an increase of poverty by less than 0.17%. Also the 1% rise in agricultural output should, according to our estimates, reduce poverty by about 1.6%. Hence, the net (compensated) elasticity should be about 1.33, well within the range that we predict for the elasticity of poverty reduction with respect to agricultural growth.

Over the period of analysis employment of unskilled workers has only increased on average by 0.44% per annum. This despite that real aggregate output rose by 7.7% per year. According to our estimates, however, the biggest obstacle for faster employment of unskilled workers is not a low responsiveness to output expansion. It is not due to ever higher real wages for the unskilled either; real wages for the unskilled increased by only 2.4% per annum compared to 3.6% for skilled workers and 1.8% for capital. In fact, this moderate wage increase only causes a very small negative impact on unskilled labour demand compared to the very dramatic increase that expanding output caused. The real source for slow unskilled job creation is the enormous bias of technological change against unskilled workers. This also constitutes a significant obstacle to a faster rate of poverty reduction. This result thus underlines another cost associated with the low priority given to investments in research, development and especially technological adaptation in Chile.

Data Appendix

The measures of agricultural output and total output, both at the national level, as well as the regional level come from the National Accounts maintained by Chile's Banco Central. We used the accounts with constant figures with 1986 as base year. We additionally developed an indicator of agricultural processing value added. The development of this series required two main data sources, the Chilean Economy's input-output matrix (1996), and the yearly "Encuesta Nacional de la Industria Anual" ENIA (1980-1998). (The Input-Output matrix is available in Banco Central (1998), ENIA surveys performed by the Instituto Nacional de Estadísticas (INE), is available in print at their public library.) With the aid of the input output matrix we first identified the industrial sectors that use agricultural output as their main input. Example of these sectors include: meat processing, milk processing, and wine industries. With the aid of the ENIA surveys we determined the share of total industrial value added due to these agricultural processing industries. Finally, the product of this latter share and the national accounts industrial GDP gave us a yearly figure of agricultural processing value added both at the national level (1980-1998), and the regional level (1980-1996).

The section on employment effects of agriculture required the use of three inputs, unskilled labour, medium and high-skilled labour, and capital inputs. The monthly wages for the unskilled and skilled sectors were calculated using CASEN national surveys. The division between skilled and unskilled labour was given by the years of education of labour: workers with 8 or less years of education were considered unskilled. The cost of capital inputs was proxied by the cost of imported capital goods given by: index of imported capital goods (Banco Central).

The section on non tradable food prices required the creation of a price index of non-tradable goods. The index was created using nominal prices of non-tradable food goods like: tomatoes, potatoes, eggs, bread, carrots, and other (monthly: 1975-2000). A Laspeyres index was constructed using weights from the national CPI (base year 1998), and was deflated by the national CPI to obtain real prices. The non-food real price index was constructed with the ratio of non-food component of the national CPI, divided by the national CPI. For the regression analysis we used yearly average of the price indexes. The real exchange rate was obtained from the Banco Central, and it is defined as the nominal exchange rate times the ratio of foreign inflation to national inflation.

For the section on poor farmers, we calculated household income for those households were the head was identified as a subsistence farmer by the CASEN national survey. Next, we separated household income between that that is generated in the farm, and that that comes from off-farm activities.

An important result is that while the economy-wide effects taking place via food prices and especially the labour market are quantitatively important the direct income effects on farmers are almost negligible.

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